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Urban Air Pollution Damage Functions: Theory and Measurement*

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The urban air pollution damage function as it is registered in differential market prices of residential property was investigated. By using somewhat more complex specifications and a set of data appearing to have greater informational content, the sensitivity of damage functions derived by means of differential land values to differences in specifications and differences in data was ascertained. If results do not appear overly sensitive to data and specification differences, then simpler and less costly specifications and data sets can be used to obtain the damage function with little or no loss in information. Everything else being equal, a control instrument more likely to approach a continuing minimization of the sum of emitter control costs and receptor damage costs can then be employed. Less obtusely stated, the question being asked here is whether any additional worthwhile information about urban air pollution damage functions can be obtained by working with property value models having somewhat more complex specifications and more complete data than in previous studies.

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Introduction

In the past five or ten years, the attention which air pollution problems have received has generated a wide variety of proposals for their alleviation. Present in all these proposals has been some notion of a damage function relating air pollution dosages to the market value of the foregone production and consumption opportunities suffered by receptors. Most academic discussions of these proposals have proceeded in a world of certainty, a world in which the values of all variables relevant to the decision process can always be consistently and costlessly predicted with zero probability of error. Most of the time property rights in the air resource are assumed to be assigned to a control agency whose tremendous clairvoyance and omniscience enables it to select effortlessly or cause to occur that level of pollutant emissions synonymous with amvient air concentrations minimizing the sum of sufferer damage and emitter control costs. Under these conditions, discussions of optimal control instruments become trivial exercises in which the only matters of interest are value-loaded statements about whose judgments are to count. Questions of real economic interest would occur only when it is supposed that the control agency knows of no systematic way to put its perfect information together.

However, once it is recognized that the discovery of that pollution program minimizing the sum of emitter and receptor costs is itself costly, important differences among alternative control instruments become apparent. Ultimately, the criterion for selection of a control instrument must be the minimization of the sum of emitter and receptor costs and all parties', including regulatory bodies', informational, contractual and policing (ICP) costs. Even assuming all property rights to the air resource are always initially vested in a control agency, the ICP costs of alternative control instruments can differ widely. For example, the use of the effluent charge as the sole instrument of control requires an intimate knowledge of the behavior of the air pollution damage function. But when the effluent charge is employed along with an already established ambient air standard, knowledge of neither the receptor damage function nor the emitter control cost function is required. All the agency need do is vary the charge in a manner causing the ambient air standard not to be exceeded. Of course, though ICP costs are relatively low in this latter case, the ambient air standard may be identified with substantial permanent or periodic deviations from minimization of the sum of receptor and emitter costs. The probable costs of this deviation must be weighed against the costs of determining the receptor and emitter cost minimizing effluent charges as well as against the costs of employing effluent charges of various degrees of suboptimuality as the sole control instrument.

The formal consideration of the preceding statements immediately leads one into the formidable quicksands of problems in decision-making under uncertainty. In this paper the intent is not to enter these sands but instead to provide an input which in the context of air pollution concerns can assist in this decision-making. The particular purpose is to investigate the urban air pollution damage function as it is registered in differential market prices of residential property. This will by no means be the first time this function has been studied in this manner [1], [2], [3, pp. 141-151], [4], [5], and [6]. By using somewhat more complex specifications and a set of data appearing to have greater informational content, this study's purpose is to ascertain the sensitivity of damage functions derived by means of differential land values to differences in specifications and differences in data. If results do not appear overly sensitive to data and specification differences, then simpler and less costly specifications and data sets can be used to obtain the damage function with little or no loss in information. Everything else being equal, a control instrument more likely to approach a continuing minimization of the sum of emitter control costs and receptor damage costs can then be employed. Less obtusely stated, the question being asked here is whether any additional worthwhile information about urban air pollution damage functions can be obtained by working with property value models having somewhat more complex specifications and more complete data than in previous studies.

The Theoretical Framework

Knowingly or unknowingly, all previous studies of the relation between air pollution and property values have used the results of Strotz [7] as a major rationale for their empirical procedures. Given certain simplifying assumptions, it is shown in [7] that if two sites are similar in all respects except air quality, the difference in their values represents the market's willingness to pay for reductions in air pollution dosages. That is, all air pollution damages will be registered in differential site values. However, if all consumers do not regard the two sites as perfect substitutes when each site has equal air pollution dosages, then some air pollution damages will be capitalized into other durable assets and losses in consumers' surplus. Nevertheless, it is shown that these sites and other assets enter individual utility functions in an additive manner. Thus differential site values can be employed to obtain a lower bound on air pollution damages, and, if the sites in question have rather homogeneous characteristics, their differential values represent a substantial portion of damages. Furthermore, when sites are not viewed as perfect substitutes by consumers, a substantial portion of damages will still be registered in differential property values, where a property includes the site and the immobile and durable improvements thereon.[1]

With one unsuccessful exception [3, pp. 141-151], previous studies of the covariation between air pollution and urban residential real estate have been strictly cross-sectional. That is, observations have been assumed to be drawn from time periods having similar initial and terminal dates and in which the effects of time's passage are presumed to be neutral. A dynamic stability has been assumed in which ". . .from any initial conditions all the variables approach their equilibrium values in the limit as time becomes infinite, i.e.,

$$\lim_{t \to \infty} x_i (t) = x_i^0$$

regardless of the initial conditions." [9, p. 261] This means that the difference between the ultimately realized or equilibrium value and the actual value of x is bounded for a sufficiently large t, and that this difference tends to zero as t becomes large. In the context of the previous studies, large has come to mean whatever single unique period for which the investigator happened to have air pollution and property observations. The expectations about air pollution dosages, on the basis of which economic adjustments to the presence of air pollution are made, are presumed to be formed solely within the single period for which information on dosages is available. In some studies the result has been that the expectations implicit in property values are assumed to have been formed on the basis of air pollution not occurring until anywhere from four to eight years later! In addition to the possibility of introducing some statistical biases,[2] this assumes what would seem an inordinate amount of prescience in the residential property market.

A more appealing hypothesis is that expectations about future air pollution dosages adapt to changes in present air pollution dosages only after some lag in time; so that if such a change has any permanent effect at all, the effect is not registered all at once over a fairly short interval of time but is instead distributed over several time periods. Specifically, temporarily assuming property value to be a function only of air pollution dosages,

$$(1) \qquad c_{it} = a + b\,E(p_{it}) + u_{it} \qquad i = 1, \ldots, n$$

where $E(p_{it})$ is the expected value of air pollution dosages at the ith property at the end of the current period, t, c is the current market value of this property, and u is an error term. All variables are expressed in logarithms throught. Expected air pollution dosages can be related to observable values by specifying an expectation generating function such as

$$(2) \qquad E(p_{it}) = p_{it-1} + \lambda_i\,[p_{it} - E(p_{it-1})].$$

This states that pollution dosage expectations are revised in proportion to the proportionate error associated with previous levels of expectations, [10, pp. 22-23].

If (2) is linear in logarithms, λ_i can be interpreted as an elasticity coefficient showing the responsiveness of current expected dosages to discrepencies between current realized dosages and the dosages expected in the previous period for the current period. In the case where the coefficient is unity, air pollution damages are simply a function of current dosages. That is, if dosages were previously expected to remain constant, implying that property values were at long-run equilibrium levels, they are now expected to remain constant at the level of current dosages. A new long-run equilibrium in the property market is therefore established. In contrast, an elasticity coefficient of zero implies that changes in current dosages have no effect upon the property market's existing long-run equilibrium. Changes in current dosages have no effect upon expected future dosages.

The expression (2) can be solved for expected dosages as a function of m ($m \geq 0$) past values of realized dosages.

$$(3) \quad E(_{it}) = \lambda_i \, [p_{it} + (1 - \lambda_i)p_{it-1} + (1 - \lambda_i)^2 p_{it-2} + \ldots + (1 - \lambda_i)^m p_{it-m}].$$

In (3), the expectation generating function can be viewed as a distributed lag with exponentially declining weights. Muth [11] has demonstrated that (3) minimizes the mean square error in forecasts of realixed dosages if the process generating these dosages is such that changes in realized dosages are first-order moving averages of random deviates. Even when the changes caused by the process are not a first-order moving average of random deviates so that mean square error of forecasts is not minimized, Cox [12] shows that (3) may still be a highly efficient predictor of realized dosages.

Making the obvious substitution of (3) into (1) and simplifying in the manner of Koyck [13], the cumulative effect upon property values of a change in realized air pollution dosages that is expected to be maintained is given by

$$(4)\quad c_{it} = a + b \sum_{t=0}^{m} \lambda_{i}^{m} p_{it} + u_{it},$$

if and only if $0 \leq \lambda < 1$. If a history of realized air pollution dosages is not well described by exponentially dealining weights, (4) can become (5) where k is some function of t.

$$(5)\quad c_{it} = a + b \sum_{t=0}^{m} \lambda_{i}^{k} p_{it} + u_{it}$$

This latter equation is consitent with any set of weights, regardless of whether older or more recent observations be given greater weight.

From the point of view of the air pollution receptor, this adaptive expectations hypothesis seems likely to have substantial validity. That is, the scene of the receptor acting as a more or less passive observer of air pollution events seems to accord with experience. He can have little knowledge of future events other than the inferences he can draw from already realized events, for an active search by him for information on future events is likely to be far too costly. This cost could be less if the planned emissions of pollution perpetrators were not entirely independent of individual receptor decisions. It would also be less if ambient air standards were imposed with some stringency or if distinct contingent claims markets in air pollution events were available for the individual receptor. The scale economies of information production and risk-spreading characteristic of these markets could make information about future air pollution events much less costly to the individual receptor. Though contingent claims markets do in fact exist for durable assets affected by air pollution, e.g., life, medical, and property insurance, the receptor must still make the costly effort to distinguish between air pollution risk's contribution to premiums and the contribution of other risks. This in spite of the high cost of making a wrong prediction about future air pollution dosages when purchasing a residential property (moving or selling is costly, particularly if done a year or two after purchase), the seemingly even higher costs of most alternative informational sources on future air pollution events make predictions based almost entirely on realized dosages economically viable. It is therefore justifiable to assume, as this study does, that the major and perhaps the only argument in the receptor air pollution expectation generating function is some history of realized air pollution dosages.[3]

For each time period dealt with in this study a long-run equilibrium is presumed to exist in all housing characteristics. Thus no supplier of housing characteristics is presumed to be able to increase his net revenues by adjusting the stock or the location of housing characteristics. Similarly, no individual buyer of housing characteristics can increase his utility by purchasing a different package of characteristics or by changing the location of his residence. In short, given the market structure, all locations are presumed to be held by those participating in activities yielding the highest property values. Our task is to explain the differences in these values. Assuming that all suppliers of housing characteristics face the same capitalization rates and cost structures in the locale of interest, these value differences are accounted for by consumer evaluations of differing locational characteristics.

The economic-theoretic basis for these consumer evaluations is the same as [1]. There the consumer was represented as having a well-behaved utility function, $U(z)$, where z is an $m \times 1$ vector of quantities of characteristics.[4] He is supposed to maximize U subject to the conditions

$$\begin{aligned} c'x &= y \\ z &= Bx \\ z, x &\geq 0, \end{aligned}$$

where x is an n x 1 vector of quantities of the goods in which the characteristics are embodied, c is the n x 1 vector of prices associated with these goods, B is an m x n matrix, and y is income. The utility maximizing quantities of the various characteristics are functions of the prices of the various goods in which they are embodied and consumer incomes. As [1] shows, it is possible to write $c = c(z, y)$, where $c(\cdot)$ is the consumer's offer function. However, this unrestricted formulation could require that the consumer evaluate all goods in which a characteristic entering his utility function is embodied. Since a house may embody characteristics found in a wide variety of other goods, a complete empirical specification of this offer function would require the simultaneous specification of offer functions for all those goods also embodying some of the characteristics embodied in housing.

Actually, as Lancaster [14] emphasizes, the characteristics embodied in some goods, for given relative prices of goods, will represent inefficient consumer choices and can therefore be disregarded. But in any empirical application the elements of the B matrix must be specified, an ambiguous task for which there is little if any theoretical guide. Neither the researcher nor the consumer are likely to possess the skills and willingness to survey and count completely and simultaneously all characteristics to be found in housing. If one introduces, as we shall, dated characteristics and stochastic characteristics, the mental feats required for the researcher to perceive all gradients of the consumption technology are indeed astounding. The same point can be made for the consumer. Though it is true a great many things can be measured if the urge for measurement is strong enough, the very fact that some urging is required implies a cost. One can therefore reasonably presume the consumer strives to satisfy a craving for simplicity by means of some aggregation procedure reducing at some stage the rank of the B matrix he must confront.

Consider the following scheme which in certain of its fundamentals is similar to that presented by Green [15, pp. 9-32]. In order to conserve his means of surveying and counting, the consumer completely disregards some commodities and aggregates over others having collections of characteristics intended to serve a more or less broad function, e.g., housing, food, entertainment. In the first decision stage, the allocation of the consumer's budget on the basis of price indices of the commodities forming such broad groupings does not make an intolerable difference in the decision problem with which the consumer is confronted. We assume his demand for the broad grouping "housing" or any other broad grouping to be altogether independent of air pollution considerations. Only after this first-stage decision has been made are individual collections of housing characteristics evaluated. One of these evaluated characteristics is the presence of air pollution.

It is well known that if the utility function U[z(x)] is strongly separable in the sense of Goldman and Uzawa [16], the quantity of each commodity purchased with the consumer operating under a two-stage procedure will be identical to the quantity purchased if the consumer's purchases had all been made in one stage. That is, if we assume the commodities embodying characteristics entering the utility function are divided into three groups H, I, and J, where H is housing, then strong separability is defined as:

$$\frac{\partial\left(\frac{U_i}{U_j}\right)}{\partial x_h} = 0 \qquad \text{for } h \in H,\; i \in I,\; j \in J,\; H \cap I \cap J = \emptyset$$

where i, j, and h are commodities having at least one characteristic not in common. Given this restriction upon the utility function, if the consumer allocates his budget in the first stage so as to maximize a utility function having groups of characteristics embodied in groups of goods as its arguments, then the consumer can maximize his overall utility by maximizing a utility function in the second stage having characteristics embodied in individual goods as its arguments. In our context, a utility maximizing allocation of the consumer's budget among housing locations is therefore not inconsistent with a utility maximizing allocation of this budget among all goods. This result does not necessarily imply that groups of characteristics interdependent in demand do not exist. It only implies that the individual property buyer groups all interdependent characteristics so that he makes his final house purchase decision in terms of the mutually independent groupings which we are presumed to be able to describe. Conceptually, these separability notions imply that air pollution damages to housing will be relatively greater because they permit the consumer to respond to changes in air pollution dosages only by a redistribution of his expenditures among characteristics in the housing group rather than among all groups. However, in terms of estimating the effects of air pollution upon the demand for housing, the failure of these separability notions to be fulfilled means that the characteristics of certain goods which contribute to the explanation of variations in offer prices for housing are not being included. If these characteristics are not orthogonal to air pollution dosages and if they contribute positively to offer prices, then air pollution damages will tend to be biased downward. Similarly, if they contribute negatively, air pollution damage estimates will tend to be biased upward.

On the basis of the above statements, the offer function to be estimated is a quantity weighted average of implicit prices for characteristics embodied in a commodity making up something called a housing location. It should be emphasized that each price observation is presumed to be attached to one unit of a unique commodity, a housing location. The c_i are therefore observations on the market prices of distinct goods. They represent T observable prices of T different goods. Each refers to one and only one house.[5] Therefore, making the not unreasonable assumption that each residential property owner owns only one house and that each residential property is owned by only one owner, the offer price for housing and consumer expenditures for housing are identical.[6]

To this point, the discussion has implicitly assumed that for those variables whose future values are not largely subject to the property owner's control, only the expected value of the variable's frequency distribution is relevant to the owner's decisions. In other words, certainty equivalence has been assumed such that the first moment of the air pollution frequency distribution at a particular site adequately summarizes the entire distribution. The disutility the receptor attaches to future air pollution dosages has been treated as identical in situations where perfect foresight is lacking and in situations where it is present. The receptor's behavior at a point in time is therefore viewed as invariant with respect to the probability of error in his forecast of expected air pollution dosages.[7]

In the following, we continue to express the multivalued expected results associated with alternative time-states as probability distributions rather than assuming that the property buyer sets these states out in extensive form.[8] That is, we add to the set of air pollution characteristics relevant to the property buyer various derivative statistical measures of air pollution frequency distributions. In particular, by assuming U(z) to be cubic in reduced air pollution dosages, we show the relevance of the first three moments of the air pollution frequency distribution to air pollution damages.[9] However, this demonstration does not assert that only the first three moments count, for, as Richter [22] has shown, an nth degree polynomial for U(z) implies and is implied by the assumption that only the first n moments are to count. In short, by assuming a utility function of the proper degree one can justify the use of any number of moments (until their values tend to zero).

For a given time interval, let p be the difference per unit time between the greatest concentration of air pollution in an urban area and the pollution dosages occurring per unit time at any particular location within the same urban area. Assuming as before that fewer air pollution dosages are preferred to more, we can write

$$(6) \qquad U = U(p), \text{ and } \frac{\partial U}{\partial p} > 0 \text{ uniformly.}$$

Expanding (6) in a Taylor series about a mean value, $\bar{p}$, ignoring moments above the third, and taking the expected value yields

$$(7) \qquad E[U(p^*)] = U(\bar{p}) + \left(\frac{\sigma^2}{2}\right)\left(\frac{\partial^2 U}{\partial \bar{p}^2}\right) + \left(\frac{\sigma^3}{6}\right)\left(\frac{\partial^3 U}{\partial \bar{p}^3}\right)$$

As a random variable, $p = p^*$, and

$$\sigma^2 = E(p^* - \bar{p})^2$$

$$\sigma^3 = E(p^* - \bar{p})^3$$

As noted above, taking the utility of cleaner air to be a function of the first three moments of p*'s distribution about its mean is equivalent to assuming U(p*) to be a cubic of the form

$$(8) \qquad U(p^*) = p^* + bp^{*2} + gp^{*3}$$

Upon taking the expected value of (8) we obtain

$$(9) \qquad E[U(p^*)] = E(p^*) + bE(p^*)^2 + g\ E(p^*)^3$$

where

$$(10) \qquad E(p^*)^2 = \sigma^2 + [E(p^*)]^2$$

and

$$(11) \qquad E(p^*)^3 = \sigma^3 - 2[E(p^*)]^3 + 3E(p^*)^2\ E(p^*).$$

Substitution of (10) and (11) into (9) gives

$$(12) \qquad E[U(p^*)] = E(p^*) + b[E(p^*)]^2 + g[E(p^*)]^3 + [3gE(p^*) + b]\sigma^2 + g\sigma^3$$

Taking the partial derivative of (12) with respect to $E(p^*)$, we get

$$\frac{\partial E[U(p^*)]}{\partial E(p^*)} = 1 + 2bE(p^*) + 3g[E(p^*)]^2 + 3g\ \sigma^2.$$

$$= 1 + 2bE(p^*) + 3g[E(p^*)]^2 + 3g(E(p^*)^2 - [E(p^*)]^2).$$

$$(13) \qquad = 1 + 2bE(p^*) + 3g\ E(p^*)^2.$$

This is a quadratic having roots

$$\frac{-2b \pm [2b^2 - 4(3g)]^{\frac{1}{2}}}{6g}$$

Thus if the marginal utility of cleaner air is always to be positive, then

$$2b^2 < 4(3g).$$

That is, for (13) to be positive, it is required that

$$(14) \qquad b^2 < 3g$$

Since b^2 is always positive, g must also be positive if (14) is to hold true. Given that g must be positive, the sign for $\frac{\partial^2 U}{\partial \bar{p}^2}$ comes from (7) and (12) where

$$(15) \qquad \frac{\partial E[U(p^*)]}{\partial \sigma^3} = \frac{\frac{\partial 2U}{\partial \bar{p}^2}}{2} = 3g\ E(p^*) + b = 3g\ \bar{p} + b.$$

Clearly, if $\bar{p} \leq \frac{-b}{3g}$, then $\frac{\partial^2 U}{\partial \bar{p}^2} \geq 0$. The assumption of $\bar{p} \leq \frac{-b}{3g}$ amounts to an assumption of risk averse behavior on the receptor's part. The correct sign for $\frac{\partial^3 U}{\partial \bar{p}^3}$ can also be obtained from (7) and (12) since

$$(16) \qquad \frac{\partial E[U(p^*)]}{\partial \sigma^3} = \frac{\frac{\partial^3 U}{\partial \bar{p}^3}}{6} = g$$

The requirement from (14) that g be positive therefore assures that $\partial^3 U/\partial \bar{p}^3 > 0$.

Our conclusion from (15) and (16) above is that the receptor attaches disutility to increasing uncertainty, as measured by variance, of air pollution dosages. Furthermore, given equivalent expected values and variances, he prefers those air pollution frequency distributions skewed toward the lower range of dosages. Treating these derivative statistical measures as distinct characteristics, we therefore view each residential property as giving rise to at least three air pollution characteristics, the mean, the variance, and the skewness of the property's air pollution frequency distribution.

As we earlier noted, there is no particularly defensible reason other than investigative convenience to assert that only the first moment or, for that matter, the first three moments of the distribution of air pollution dosages enter as arguments into the housing consumer's utility function. In effect, the problem of discovering what these arguments are with respect to air pollution remains an empirical question. The best the investigator can therefore hope to do is to specify models capable of being empirically evaluated.

There exist several alternative decision criteria conceivably applicable to a sufferer faced with uncertainty about the future pollution dosages to which he will be subject. Each of these alternatives implies a preference ordering over possible future states based upon a criterion other than the maximization of expected utility.[10] Many of these criteria cannot be meaningfully described in probability terms. One often cited criterion of this sort is the minimax decision criterion of Wald [24] where (in our context) the sufferer is presumed to know with zero probability of error that future air pollution dosages will fall within a given range. The sufferer's objective is then to make his maximum possible loss as small as possible, where the loss is defined in terms of the discrepancy between the realized consequence of an act and the utility which would have been obtained if the future state of nature had been correctly forecast. As Hildreth [25] notes, the common sense appeal of a criterion of this sort depends directly on how small this maximum possible loss can be made. If this maximum possible loss is likely to be small relative to the costs of acquiring that information necessary for more accurate forecasts -- as it is likely to be for air pollution situations even when increased risks of health effects are accounted for -- then the criterion would appear to be reasonable. If, for example, with wide fluctuations in air pollution dosages and damages a sufferer employing this criterion can come within $100 of his minimum possible losses under perfect certainty, then the criterion might arouse the sufferer's interest. But if the discrepency results in his death or even damages of $1000 he might display no interest whatsoever.

The article of Hildreth [25] referred to above derives an expression from a Cobb-Douglas production function for a firm's demand for an input where thr firm employs a minimax decision criterion. A somewhat more general development applicable to the minimax decision criterion of an air pollution sufferer can be obtained with the aid of some easily interpreted diagrammatics.

Assume that the sufferer acts as if air pollution dosages over whatever time period is deemed relevant are bounded above and below by fixed values. Let $U(p^*)$ in Figure 1 represent the upper bound of these dosages and $U(p_*)$ the lower bound. The sufferer would act as if any dosage, p, in the interval $\{U(p^*), U(p_*)\}$ could occur. Since $\frac{\partial U}{\partial p} < 0$, movement up the U axis represents increasing utility levels and decreasing air pollution dosages. From the sufferer's point of view, any reduction whatsoever in dosages is gratis. Similarly, any increase in dosages is presumed to be utterly independent of any activities the sufferer might undertake. Thus on the U axis, each utility level is associated with a unique dosage of air pollutants over whose presence or absence the sufferer knows he can exercise no influence.

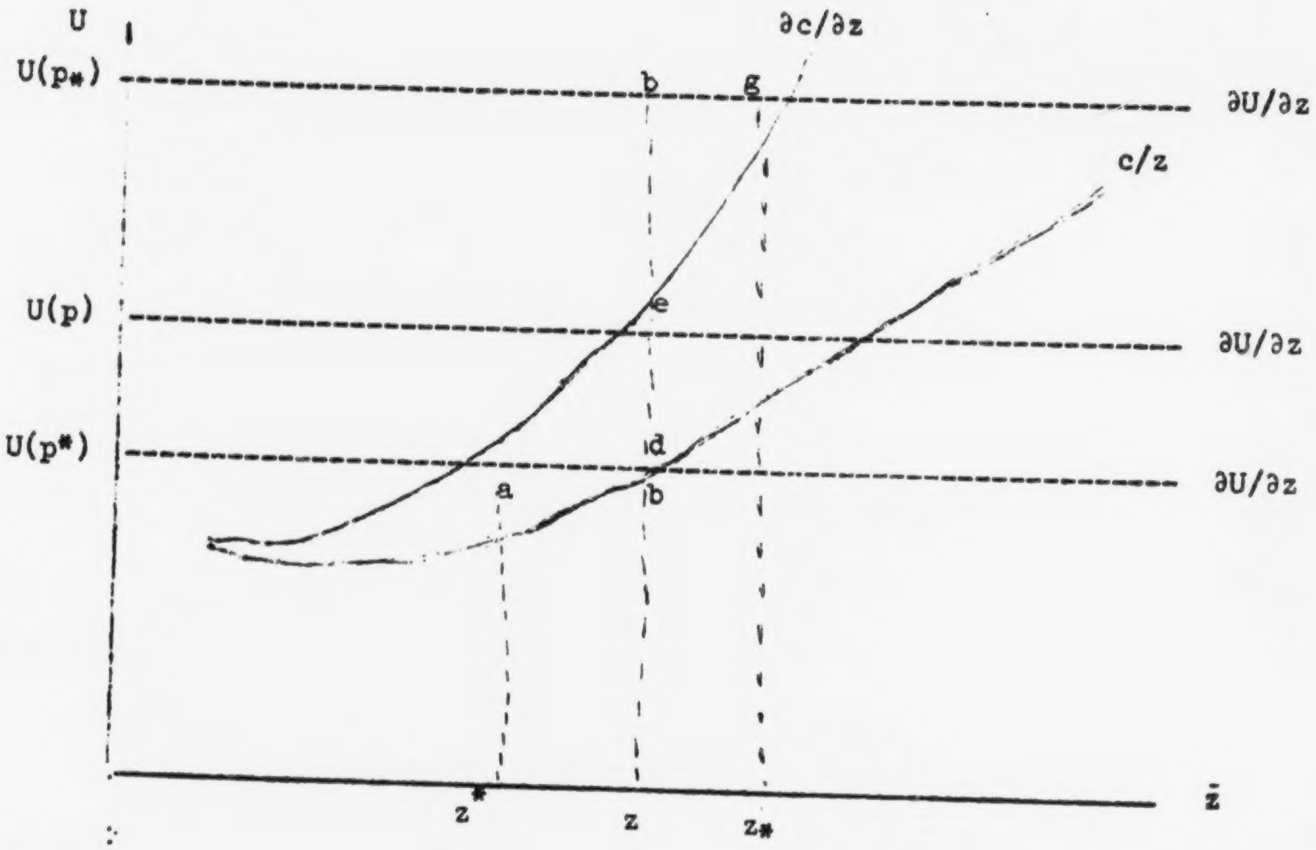

Figure 1. Possible Sufferer Losses from Errors in Forecasting Pollution Dosages

With any given level of air pollution dosages, there can be combined various quantities of a bundle, z, of housing characteristics.[11] The additional utility to be obtained at a given dosage level with the consumption of an additional unit of the bundle of housing characteristics is given by $\frac{\partial U}{\partial z}$. The marginal costs and average costs of obtaining an additional unit of the bundle of housing characteristics are given respectively by $\partial c/\partial z$ and c/z. Utility is assumed to be homogeneous of degree one in dollars. The increasing $\frac{\partial U}{\partial z}$ associated with decreasing pollution dosages permits pollution to affect the utility obtainable from the elements of the bundle of housing characteristics. The sufferer's problem is to select that quantity of the bundle of housing characteristics minimizing his maximum possible loss from air pollution dosages.

Temporarily assume the sufferer acts as if the pollution dosage is to be p, where p lies somewhere between p^* and p_*. He therefore chooses to consume z units of the bundle of housing characteristics, where z is someplace in the interval $\{z^*, z_*\}$. If p^* rather than p is ultimately realized, the sufferer has consumed $ab = z^*z$ units of the bundle of housing characteristics whose cost outweighs the utility these units provide him. His total loss is represented by the area adb. On the other hand, if p_* rather than p is ultimately realized, the sufferer has failed to consume $fg = zz_*$ units of the bundle of housing characteristics whose cost does not outweigh the utility these units could provide him. The loss he suffers by overestimating future pollution dosages is given by the area efg. Clearly, the minimax rule amounts to finding a p such that the quantity of the bundle of housing characteristics associated with that p implies adb = efg. Thus unless the marginal costs of consuming additional units of the bundle of housing characteristics are constant, the minimax pollution dosage to expect is a function of the maximum and minimum expected dosages and the costs of obtaining additional units of the bundle of housing characteristics.[12]

To this point in this essay, little has been said about the relations among utility, air pollution dosages, air pollution damages, and the market prices of housing or housing sites.[13] The connection between air pollution and land and property prices, and between land and property prices and air pollution damages has been relatively neglected. Accepting the existence of an inverse relation between air pollution dosages and these economic measures still leaves a great many interesting questions unanswered and untested. Three of these questions are of particular interest here.[14] First, given that some insight can be acquired into the form of the relation between property values and air pollution doasges, are there any inferences which can be drawn from this insight about the form of the relation between air pollution dosages and air pollution damages apart from any capitalization of the latter into property values? Second, what is the likely form of the relation between air pollution dosages and property values? In an earlier effort of the writer and R. J. Anderson, Jr. [1, p. 25], empirical results over some interval of increasing air pollution dosages pointed to the possibility of a declining marginal property damage function as well as a declining property damage elasticity. No explanation of the phenomenon was offered however. Finally, what proportion of the damages to a property are capitalized into land or site values relative to improvement values? Does the elasticity of land values differ from the elasticity of improvement values with respect to air pollution dosages?

In the following, the household is viewed as a firm selling a bundle of housing characteristics to itself for a price equal to its maximum willingness to pay for that bundle. For simplicity, all households are treated as identical. In order to remain consistent with our previous derivation of the offer function, these households are viewed as purchasing various goods to obtain this bundle of characteristics. The bundle of housing characteristics and the collection of goods used to obtain these characteristics are assumed to fulfill the conditions for a composite good. The function relating the bundle to the collection of goods is homogeneous of degree one. All market prices are competitively determined. The collection of goods embodying housing characteristics can be varied over a given state of nature.

Given the above conditions, the householder's problem is to maximize

$$U = cz - c_x x \tag{17}$$

subject to

$$z = z(x),$$

where, as before, U is the householder's total utility, z is the bundle of housing characteristics, and x is the collection of goods. The unit cost of the collection of goods is c_x and the unit cost of the bundle of characteristics is c. U can thus be regarded as a measure of the Marshallian consumer surplus. The first order condition for a maximum is

$$\frac{\partial U}{\partial x} = c\,\frac{\partial z}{\partial x} - c_x = 0. \tag{18}$$

Air pollution dosages are specific to location. For locational equilibrium to prevail, it is necessary and sufficient that each household occupy the most expensive available property it is willing to buy and that each available property be occupied by that household willing to pay the highest price. The change in the bundle of characteristics resulting from any change in the household's location is given by

$$dz = \frac{\partial z}{\partial x}\,dx. \tag{19}$$

The condition of (18) implies

$$c_x = c\frac{\partial z}{\partial x}, \tag{20}$$

and the change in consumer surplus accompanying such a change in location is

$$dU = (z\ dc - x\ dc_x) + (c\ dz - c_x\ dx). \tag{21}$$

The first-order condition of (20) is always fulfilled. Therefore, upon substituting (19) and (20) into (21) we have

(22) $$dU = (z\ dc - x\ dc_x) + (c \frac{\partial z}{\partial x} dx - c\frac{\partial z}{\partial x} dx) = dc_x - x\ dc_x.$$

If there is to be locational equilibrium, dU = 0. Thus (22) becomes

(23) $$dc_x = \frac{z}{x} dc.$$

The householder's willingness to pay for a bundle of housing characteristics at a particular location is viewed as a function of the magnitude of air pollution dosages at that location. That is, $c = c(p)$, where, as before, p refers to air pollution dosages. By hypothesis, $\frac{dc}{dp} < 0$. Therefore, from (23)

(24) $$\frac{dc_x}{dp} = (\frac{z}{x}) \frac{dc}{dp} < 0.$$

The ratio of marginal damages for goods embodying housing characteristics to marginal damages for the bundle of housing characteristics is obviously dependent upon the ratio of z to x. Furthermore,

(25) $$\frac{d^2c_x}{dp^2} = (\frac{z}{x}) \frac{d^2c}{dp^2} + \frac{dc}{dp} \frac{d(z/x)}{dp}$$

As noted above, $\frac{dc}{dp}$ is negative. One of the major economic features of air pollution is that it reduces the output to be obtained from any given mix and magnitude of inputs. Thus $d(z/x)/dp$ is also negative. Since a positive quantity of characteristics can be obtained only with a positive quantity of goods, z/x is positive. The slope of the marginal air pollution damage function for goods, d^2c_x/dp^2, is therefore easily seen to be dependent upon the sign of the slope of the marginal air pollution damage function for housing, d^2c/dp^2. If the latter proves to be negative so that air pollution damages to housing are increasing at a decreasing rate with respect to increasing air pollution dosages, then, when viewed in the aggregate, air pollution damages to those goods capable of providing housing characteristics must also be increasing at a decreasing rate. The converse is of course true when d^2c/dp^2 is found to be positive.

In location theory, it is widely acknowledged that d^2c is negative with respect to any variable which differs in magnitude from one location to another and to whose increased presence disutility is attached. The reasoning goes as follows. Locational equilibrium requires that households be unable to increase $(U = cz - c_x x)$ by any move. Given that air pollution and property values are inversely related, the household can increase the quantity it can purchase of the goods embodying any bundle of housing characteristics for a particular money outlay by moving to an area subject to greater pollution dosages. However, the household's willingness to pay for a given bundle of housing characteristics at the higher pollution location decreases. This implies that in equilibrium

$$(26) \qquad \frac{\partial c_x}{\partial p} = \frac{\partial c}{\partial p} .$$

In short, in locational equilibrium a change due to pollution in the price of goods embodying a given bundle of housing characteristics is exactly offset by the change in the householder's willingness to pay for this bundle. Furthermore, if this equilibrium is to be maintained, it is necessary that with increased air pollution householder willingness to pay decline at a rate more rapid than the decline in the market prices of goods embodying housing characteristics. Otherwise the location of each householder would be indeterminant or each householder could always increase U by moving to locations having the highest air pollution dosages. This result is fully consistent with (25), given that d^2c/dp^2 is negative.

The above implies that those households for whom $\frac{\partial c}{\partial p}$ is relatively high will tend to locate where air pollution dosages are relatively low, and those households for whom $\frac{\partial c}{\partial p}$ is relatively low will tend to locate where such dosages are relatively high. The equilibrium condition of (26) states that when $\frac{\partial c}{\partial p}$ is high $\frac{\partial c_x}{\partial p}$ must also be high. Pollution sensitive people bid the most for pollution free locations because these people attach greater disutility to a little bit of pollution than does anyone else. Sites are thus assigned on the basis of the individual's pollution sensitivity, with the most pollution sensitive being assigned to the low pollution sites and the least pollution sensitive being assigned to the high pollution sites. The absolute magnitude of $\frac{\partial c}{\partial p}$ thus declines over an array of locations ordered on the basis of the magnitude of the air pollution dosages to which they are subject. One would therefore expect to find when comparing marginal property damage functions for high pollution locations to those for low pollution locations that the absolute magnitude of the latter is less than the former. If the "production function" $z(x)$ is homogeneous of degree one so that changes in x lead to identical percentage changes in z, then this conclusion also holds for the elasticity of the value of a bundle of housing characteristics with respect to air pollution dosages.

Additional interesting results can be obtained by retaining the same assumptions but revising the simple formulation beginning with (17) so that land and nonland goods embodying housing characteristics are distinguished. Land goods include all permanent features of a location. For practical purposes, they can include such features as utilities and man-made changes in topography as well as the unimproved space associated with the site. With this formulation, (17) becomes

(27) $$U = c\ z(x_L\ ,\ x_N) - c_L\ x_L - c_N\ x_N,$$

where U, c, and z are defined as before. x_L and x_N indicate land and nonland goods respectively and c_L and c_N are their respective prices. The first order conditions for a maximum are

(28) $$\frac{\partial U}{\partial x_L} = c\,\frac{\partial z}{\partial x_L} - c_L = 0$$

$$\frac{\partial U}{\partial x_N} = c\,\frac{\partial z}{\partial x_N} - c_N = 0.$$

Performing the same set of operations as in (19)-(23), one obtains

(29) $$z\ dc - x_L\ dc_L - x_N\ dc_N = 0$$

and

(30) $$\frac{dc_L}{dc_N} = \frac{zdc}{x_L} - \frac{x_N}{x_L}\ .$$

Thus, in value terms, given that $z \geq x_L + x_N$ and $dc > 1$, it follows that (30) will be greater than unity. That is, marginal air pollution damages to a bundle of housing characteristics will be registered to a greater extent in land than in the nonland goods which contribute to housing characteristics. In general, the less the marginal value product of land in terms of the value of the characteristics it produces, the greater will be the dc_L caused by a change in air pollution dosages. Again, if the "production function" $z(x_L, x_N)$ is homogeneous of degree one so that eqiproportional increases in x_L and x_N lead to identical percentage increases in z, then marginal rates of substitution between land and nonland goods remain unchanged and a conclusion similar to (30) holds for the elasticity of land values and nonland values with respect to variations in air pollution dosages. If the elasticity of substitution between land and nonland goods is positive and if the supply elasticity of land goods is less than the supply elasticity of nonland goods, the conclusion of (30) is strengthened, for the magnitude of dc_L relative to variations in dc will be accentuated. However, as land goods are substituted for nonland goods because of the change in relative prices, nonland goods will exhibit the greatest relative decline in quantity employed as a consequence of an increase in air pollution dosages.[15] In other words, given the correctness of the preceding assumptions, for a given population air pollution increases what is typically known as "urban sprawl".

A number of reasonable offhand estimates about the likely real magnitudes and signs of the above parameters can be made. First, land goods probably contribute less than one-third of the value of a typical bundle of housing characteristics, if that much. Second, the elasticity of substitution of land for nonland goods in housing is generally acknowledged to be slightly less than unity, e.g. Muth [27, p. 82]. Finally, it is intuitively obvious that the supply elasticity of residential land in an urban area must be substantially less than the supply elasticity for nonland residential goods. There are thus strong a priori empirical as well as theoretical grounds for expecting land values to be more responsive than nonland values to variations in air pollution dosages whenever any statistical tests are made. Contrary test results would more likely justify dismissal of the test on grounds of specification or measurement error rather than a failure to accept the hypothesis.

The Empirical Background

The city from which the data used in this essay come is Chicago, Illinois, a city possessing the distinction of having the dirtiest air of any large city in the United States [31]. Information on monthly ambient air concentrations of suspended particulates and sulfur dioxide was collected for each month from January, 1964, through December, 1967. Information on annual dustfall was obtained for each year from 1964 through 1967. This air pollution data was combined with data on 1288 individual single family residential property transactions whose mortgages had been insured by the Federal Housing Administration during the same period.[16] Excluding transformations, 297 separate bits of information on the air pollution history and the transaction, buyer, improvement, site, and neighborhood characteristics were collected for each individual transaction. Air pollution histories were obtained from the Chicago Air Pollution Control District, while information on characteristics was procured from the Federal Housing Administration and various secondary sources.

The completely disaggregated and highly detailed nature of the above data permits the testing of nearly all the hypotheses set forth in the previous section. In addition, this data allows one to acquire insight into whether the several possible sources of statistical bias in previous studies are indeed actual.

With one unsuccessful exception [3, pp. 141-151], previous studies of the covariation between air pollution dosages and residential property values have employed census tracts as the unit of observation exclusively. Tract boundaries usually contain within them at least several hundred housing units. These areas, according to the compilers of the census [32, p.1], are supposed to be relatively uniform with respect to population characteristics, economic status, and living conditions, so that measures of central tendency for the individual tracts presumably characterize closely all units in the tract. In the empirical analysis, group averages of residential unit values are customarily regressed upon group averages of the residents,' residential units', and tract's characteristics. The aggregation thus precedes analysis and the researcher therefore has had no basis upon which to evaluate the tolerability of any bias which the aggregation may introduce. Of necessity, it has been assumed that a knowledge of the aggregate relation (the function relating the aggregates) leads to the same value of the dependent variable (offer price) as does a knowledge of the disaggregated relation.

Aggregation bias can arise because the assumption of considerable homogeneity among the residents and residential locations within a single census tract may in fact not be fulfilled, and because the air pollution dosages to which each residential location is subjected may not be linearly related to whatever measure of "average" tract pollution dosages is employed. Assume the disaggregated offer function for bundles of housing characteristics is given by

$$c_{ij} = a_i + b_i p_{ij} + u_{ij} \qquad (1) \qquad \begin{array}{l} i = 1,\ldots,n \\ j = 1,\ldots,m \\ n > m \end{array}$$

where i refers to residential location, j to census tracts, a and b are coefficients to be estimated, and u is an error term having the customary properties. For simplicity, we assume that air pollution, p, is the only argument of offer price, c, which differs across residential locations.

Average tract offer prices and pollution dosages are clearly

$$C_j = \frac{\sum_{i=1}^{n} c_{ij}}{n}, \text{ and } P_j = \frac{\sum_{i=1}^{n} p_{ij}}{n} \tag{2}$$

The aggregate relation is therefore

$$C_j = a + bP_j + u_j \tag{3}$$

In other words, the average offer price on a census tract depends on the pollution dosages suffered by the tract's n residential locations. This seems a reasonable enough procedure, since (3) continues to be linear and includes an error term whose expected value remains zero for all P_j provided that

$$a = \frac{\sum_{i=1}^{n} a_i}{n}, \text{ and } u_j = \frac{\sum_{i=1}^{m} u_{ij}}{n}. \tag{4}$$

However, according to Green [15, pp. 99-103], an aggregation bias can be introduced because E(b), where E is the expectations operator, is a weighted mean of the b_i. Since the weight can be shown to be equal to the coefficient of a simple regression of P_j on p_{ij}, the more pollution sensitive households whose dosages have been subject to greater variability than other households in the census tract would bias upward the estimated aggregate pollution damage coefficient, $\hat{b}$, relative to the true estimate, $\hat{b}_i$. The reason is simply because the weight registered for these households is greater than their mean pollution dosages. Similar conclusions hold for pollution insensitive households subject to lesser relative variations in pollution dosages, except that in this case b would be biased downward relative to b_i. Available data permit the estimation of seperate offer functions for Chicago using census data and using the disaggregated FHA data.

The simultaneous availability of the census and disaggregated FHA data permit the investigation of two other possible sources of bias in previous studies. First, in all these previous studies, observations on air pollution dosages have post-dated observations on offer prices as well as most other explanatory variables by anywhere from four to seven years. In [1, p. 13], it is shown under fairly weak conditions that such discrepencies can cause the negative influence of air pollution dosages to be understated. Clearly, there is no reason to think that the air pollutant levels to which property owners in year t are responding necessarily correspond to the air pollution levels prevailing in t + k (k > 0). There is no reason to think that the owners at a given location were the same in the two sets of years, nor is there any reason why the air pollution levels must be the same. At a particular location, the magnitude of the response to changes over time in air pollution levels need not be exactly proportionate to the change in pollution levels. Furthermore, those locations which prove to be high air pollution areas in t + k may have experienced relatively low levels up through t. That is, the relatively high and relatively low air pollution locations may have differed in t + k and t. Obviously, the quality of statistical resolution one obtains with all these possibilities of measurement error leave cause for scepticism about the meaningfulness of results. The data available to the present study allow one to ascertain whether such scepticism is justified, since property value data for any one of four years can be combined with air pollution data for any one of four years.

A perhaps somewhat more mundane advantage of the present data is that it employs the actual market sales prices of each bundle of housing characteristics. Except for rental properties, the offer prices set forth in the census volumes are defined as owners' estimates of what their properties would sell for if offered for sale at the time of the census [32, p. 7]. Response errors to a question of this sort could be large, though it is generally agreed that the errors are random and on the average tend to be offsetting. It is therefore thought that the traditional errors-in-variables specification yields reliable estimates, particularly when some prior aggregation occurs on the basis of census tracts. Nevertheless, the only available evidence to support this hypothesis is a solitary article [33] published in 1954 which showed that owners' estimates of offer prices in the 1950 census exhibited random variation around the "true" assessed values. Though the data employed in the present study can neither confirm nor deny this finding, it does provide the opportunity to avoid such errors as may be present in the census responses.

In spite of the extent to which previous studies have depended upon it, census data is of course not the only readily available information on the offer prices for and the components of various bundles of housing characteristics. In particular, a vast storehouse of information of this sort is readily available in the property tax assessment records of nearly every municipality in the country. Its usefulness for the problem being studied here hinges largely upon the linearity of the relation between the weights assessed valuations assign air pollution dosages and the weights actual market valuations assign them. If there is a linear relation between these weights, than assessment records could validly be employed for determining air pollution damages to residential properties. The presence in the FHA data of annual property taxes as well as actual market prices for each residential property observation permits a test of this hypothesis for the City of Chicago. Furthermore, the presence of an FHA assessed value for each observation permits a similar test between this assessed value and actual market price. In this last case, if no significant differences are apparent, one could conceivably use some combination of FHA assessed values and tax assessment records to estimate air pollution damages to residential properties.

Finally, the extremely detailed nature of the data available to this study introduces a further but rather disconnected set of advantages. First, the availability of data on residential property maintenance costs provides a basis for comparisons with the results obtained in Booz, Allen and Hamilton, Inc. [34], Michelson and Tourin [35], and similar studies. Second, previous studies because of data limitations have had to employ rather crude measures of several variables which can be expected to determine offer price. For example, median number of rooms has been used as a measure of available housing space, number of housing units per acre as a measure of lot size, and the number of housing units older than some arbitrarily specified age as a measure of the age of site imporvements. It can be shown [36, pp. 284-286] that unless the proxies are contemporaneously uncorrelated with the disturbances but correlated with the true regressors, the use of a proxy can introduce an asymptotic bias into estimates. Except for permanent resident income, the great detail in which data are available for the present study does not necessitate the use of any proxies. Finally, the available detail and number of observations allows the partition of the sample on the basis of factors which might cause differential responses to air pollution dosages. Thus one might expect and therefore test for differential responses to similar air pollution dosages by resident income, age of improvements, and neighborhood characteristics, for example.

Perhaps the major advantage offered by the detailed nature of the data is the relative (relative to that of previous studies) fineness with which air pollution dosages at each residential location could be determined. The raw data provided the principal investigator by the Chicago Air Pollution Control District consisted of monthly arithmetic means at each of twenty sampling stations for sulfur dioxide in parts per million by volume per forty-eight hours and for suspended particulates in micrograms per cubic meter per forty-eight hours. Dustfall measures were provided in terms of annual averages of monthly tons per square mile.[17] The sulfur dioxide and suspended particulate measures referred to each month seperately from January, 1964, through December, 1967, whereas the dustfall measures referred to each year from 1964 through 1967. With twenty sampling stations and twelve monthly readings, a complete set of air pollution data for the purposes of this study would have 140 items each for suspended particulates and sulfur dioxide in each year from 1964 through 1967. The actual numbers of items available are presented below in Table 1. Henceforth, dustfall will be disregarded completely, since in the empirical analysis its presence did not yield any estimates amenable to economic interpretation or statistical inference.

Table 1. Number of Distinct Air Pollution Observations in Each Year.

Year	Suspended Particulates	Sulfur Dioxide
1964	136	41
1965	140	139
1966	139	137
1967	128	126

Employing the above data, and assuming equal variances in the monthly air pollution distribution from one year to another at each of the twenty sampling stations, the hypothesis was tested by means of the "t-test" that the distributions for each year at any given sampling station were drawn from the same population. In addition, an "F-test" was applied to each set of distributions at each station to test that assumption of equal variances among distributions required for the validity of the "t-test". [38, pp. 310-319]. The exercise was performed only for those sampling stations having complete monthly records for at least two calender years in the 1964-67 period. Table 2 below shows by pairs of calender years for those stations having complete monthly records the number of stations whose means and variances exhibited statistically significant differences at the .05 levels of the two-tailed t-test and the F-test. The column labeled n indicates the number of complete record stations, t indicates the number of complete record stations not having equal means by the t-test, F is the number of complete record stations not having equal variances by the F-test, and t & F is the number of complete record stations appearing in t and F simultaneously. From Table 2 it would appear that only 1966 and 1967 exhibit statistically significant differences in dosages of suspended particulates at a substantial number of stations. However, except for 1964 and 1965, a fair number of stations have significant differences in sulfur dioxide dosages for each possible pair of years. Given that suspended particulates and sulfur dioxide have widely recognized synergic properties, the results in Table 2 provide fairly strong grounds for asserting that air pollution dosages did display substantial variation from year to year during the 1964-67 time interval. A casual inspection of the spatial distribution of residential property locations relative to the location of the air pollution sampling stations gave no cause to assert that the air pollution dosages at the residential locations behaved any differently than those at the station locations.

Table 2. Stations having Differences in Annual Air Pollution Distributions.

Years	Suspended Particulates				Sulfur Dioxide			
	n	t	f	t&F	n	t	F	t&F
1964/65	18	0	2	0	7	1	0	0
1964/66	17	0	0	0	7	4	1	1
1964/67	17	2	2	0	6	3	1	1
1965/66	19	3	0	0	18	9	2	1
1965/67	19	1	0	0	15	5	2	0
1966/67	19	11	0	0	15	6	1	0

The data on month by month suspended particulate and sulfur dioxide dosages at each sampline station were used to calculate similar dosages at each sample residential location. The calculations finally employed in the empirical analysis of air pollution damages were made seperately for each of the two pollutants according to the following statement:

$$(5) \quad P = \frac{\sum_{i=1}^{4} \left(\frac{\sum_{i=1}^{4} D_i}{D_i} \cdot P_i \right)}{\sum_{i=1}^{4} \left(\frac{\sum_{i=1}^{4} D_i}{D_i} \right)},$$

where P is the calculated arithmetic mean monthly dosage of suspended particulates or sulfur dioxide at a specific sample residential location, D_i

is the distance in tenths of miles from the location to the ith air pollution sampling station, and P_i is the pollution dosage registered at the ith sampling station. Stations used for this calculation were selected so as to enclose the residential location in question. No more than four stations were ever used for this purpose. P therefore amounts to a weighted index of suspended particulates or sulfur dioxide, with the wieghts consisting solely of straight line distances from the residential location to each of no less than three and no more than four sampling stations. No attempt was made to account for variations in meteorology, topography, or other factors which could contribute to differences in pollution dosages between sampling stations and residential locations.[18] The nature of the measurement error, if any, introduced by failing to account for these factors is unknown.

The Variables

As earlier noted, information was collected on an extremely wide variety of factors which are measures of and influences upon offer prices for bundles of housing characteristics. Presented below are exact definitions of those measures of offer prices and those explanatory variables which appear in the ordinary least squares regressions reported in the next section. Inclusion of variables for which information was available but which are not defined below did not ever change appreciably any of the next section's results. Furthermore, the coefficients for those variables not defined here were almost uniformly nonsignificant.[19]

Dependent variables. Seven dependent variables are to be used in this study. Unless otherwise noted, all are defined in accordance with standard FHA definitions.

COSTS is the sale price of the residential property plus those incidental costs of purchase such as closing costs which are borne by the purchaser. It is equivalent to the mortgage amount plus the mortgagor's actual cash outlay. The inclusion of these incidental costs is entirely consistent with our view of goods and generating housing characteristics. For example, the improvement in tenure security to which a costly title search contributes can be a valued characteristic of the residential property.

PSITE is the FHA-estimated market price of the site. Neighborhood characteristics, street improvements, utilities, rough grading, terracing walls, and retaining walls are all included as part of the site.

FHAVL is the FHA-estimated price that typical buyers would be warranted in paying for the (whole) property for long-term use or investment.

MAINT is the "FHA-estimated average yearly cost of maintaining the physical elements of the property to prevent acceleration of deterioration, and to assure safe and comfortable living conditions."

TAXES includes annual property taxes and any continuing nonprepayable special assessments, as estimated by the FHA. These taxes are thus the result of Chicago's assessment ratio and the property tax rate per thousand dollars of assessed valuation. In 1966, Chicago's median assessment ratio (gross assessed)value as a percentage of market value) was 35.80 percent and its median nominal property tax rate (annual tax billed as a percentage of taxable assessed value) was 5.43 percent. [43, p. 151].

MVAL6, as defined in [32, p.7], is the median of the owners' estimates on each census tract of what their properties would sell for if offered for sale in April, 1960.

MVL66 is the "average value of homes" in 1966 in each community area of Chicago. [44, pp. 151-152]. A community area is a collection of several census tracts. The writer has been unable to find any indication as to how MVL66 was established. A reasonable guess would be that it is some simple extrapolation of 1960 census figures combined with a more or less intuitive knowledge of community area trends between 1960 and 1966.

Explanatory variable The list of explanatory variables used at one point or another in the next section is long. An effort to justify in detail the use of each of these variables could require several separate essays. The justification for the use of each variable is therefore limited to a very short explanation of the sign one can expect for its coefficient and, when seemingly relevant, how the magnitude of this coefficient can be expected to vary with variations in dependent variables. A few references where more thorough explanations of the variables' behavior can be found are offered. Unless otherwise noted, all variables are defined in accordance with standard FHA definitions. Statements about the signs coefficients can be expected to assume refer to offer price as a dependent variable. When other dependent variables are relevant, explicit note will be taken. In order to facilitate reference, the explanations are presented in the alphabetical order of the acronyms assigned the variables.

ALMNM is a dummy variable having a value of one if the major living unit upon the residential property has aluminum siding and zero otherwise. This variable was employed only to explain the variation in maintenance costs. If advertisements of the aluminum industry are to be belived, aluminum's properties are such that maintenance costs are less relative to alternative types of siding.

BLK66 is the percentage of total population that was black in a community area as of April, 1966. This variable, the data source for which is [44, pp. 145-146], was used to partition the disaggregated FHA data and as an explanatory variable when MVL66 was employed as a dependent variable. As is explained in [1, pp. 17-18], the proper sign of this variable is not at all clear; nor is the meaning of the coefficient itself. Under some assumptions, a positive sign attached to a statistically significant coefficient can be taken as evidence of housing discrimination.

BRICK is a dummy variable having a value of one if a brick frame is employed to support the floors and roof of the major improvement upon a residential location and zero if not. This variable was used only when maintenance costs were the dependent variable. Since factory fabricated frames are the type of frame for which no variable is provided, it is probably the case that BRICK will take on a negative sign where MAINT is the dependent variable.

CRIMX is an index of the rate of commitment to prison of male community area residents from 1963 through 1966. The index, which was taken from Shaw and McKay [52, p. 355], is the four year rate in the community area divided by the mean rate for the Chicago metropolitan area. If the probability of criminal attacks against one's person or property increases with the absolute number of criminal offenders inhabiting one's neighborhood, then the coefficient for this variable must take on a negative sign. Or, even if criminal offenders carry out their activities beyond the neighborhood, they may have life styles making them undesirable as neighbors.

DEXPW is the distance in tenths of miles to the nearest expressway. The Chicago Skyway was not included in this calculation unless its only two access points within the City of Chicago were closer than another expressway. Expressways are presumed to reduce the time and money outlays necessary to go from one point to another. Their influence upon offer prices for residential properties is therefore expected to be positive.

DILP6 is the proportion of all residential units in a census tract classed as dilapidated in 1960. This variable, as was noted in [1, p. 15], is a composite of a neighborhood's physical appearance and the attitudes of its inhabitants about that appearance. Its influence upon offer price is clearly negative.

DISSM is the distance in tenths of miles to the intersection of State and Madison Streets in the Loop area of downtown Chicago. Given that the disamenities of residential locations near the center of the city don't outweigh the savings in travel costs, the coefficient for the variable must be negative.

DLAKE represents the distance in tenths of miles of the residential location from Lake Michigan. Since this lake is widely acknowledged to be one of Chicago's prime esthetic and climatic amenities, the coefficient for this variable will be negative.

DTRAN is the distance in tenths of miles from the nearest private or public rail transit facility. Again, the expectation of travel savings outweighing disamenities leads to a negative value for this variable.

DTTYP is a dummy variable having a value of one if the major living unit was completely detached from any other living unit and zero otherwise. The sign for the coefficient for this variable is expected to be positive, given that a nondetached structure offers less privacy and requires agreement with adjoining owners before modifications of many exterior characteristics can be undertaken.

FRAME is a dummy variable having a value of one if a wooden frame is employed to support the floors and the roof and zero if not. Given that wooden frames typically cost more to maintain and have shorter expected lives than most other frames, the coefficient for this variable is expected to be negative in sign.

HSAGE is the age in calendar years of the major living unit at the residential location. It was calculated as the difference between the year of sale and the year in which the major living unit was built. Because of depreciation and obsolescence, the coefficient for this variable is expected to be negative. It is generally acknowledged by most writers on housing, e.g., Grigsby [45], that rates of depreciation and obsolescence tend to decline with age.

INCID is what the FHA terms incidental costs. These are said to include the FHA examination fee, service charges, costs of title search, charges for the preparation of deed and mortgage documents, recording fees, and similar items. They do not include deposits for unaccrued taxes, insurance premiums, and similar items. Clearly, since COSTS includes INCID, and since INCID is non-negative, its sign must be positive.

INCM6 is the median family income on a census tract in April, 1960. Given that bundles of housing characteristics are not inferior goods, the sign of the coefficient for this variable must be positive. Since each observation of this variable is a measure of central tendency, it is probably a fair measure of permanent income.

INC66 is the "average income" in April, 1966, in each residential location's community area. Since a community area contains within it several census tracts, it is more likely to be made up of a broad and heterogeneous collection of occupational groups who, when considered together, have transitory incomes averaging out to zero. This measure of income, given that it does not contain other sources of measurement error, should thus be better than INCM6 as a measure of permanent income.

INDUT is the hundreths of square miles devoted to industrial uses within the square mile whose center is the residential location. Data for this variable was obtained from maps published in the several volumes of [54]. The variable is intended to account for the presumed disamenities of noise, unsightliness, etc., of residing adjacent to or within an industrial area. Dirt disamenities are registered in the air pollution variable. If expectations of a shift of residential uses to industrial uses are not predominant, this variable has a depressing effect upon offer price.

LIVAR is "the total square foot area of a house appropriately improved for the intended use and in compliance with the mimum property standards for new homes and with generally accepted criteria for existing homes. It includes family rooms, improved recreation and attic rooms, cantilevered overhang of rooms and practically all improved areas in the house." Additional living space permits a greater variety of activities and houses a greater number of people. The sign of the variable's coefficient can therefore be expected to be positive.

LOTSZ is simply the number of square feet in the lot. Since greater lot size permits a greater variety of positively valued activities to be undertaken, the covariation between offer price and lot size should be positive. Alonso [26] and Muth [27] devote substantial energies to the theoretical significance of this variable in markets for bundles of housing characteristics. No information was available for this variable in the 1964 disaggregated FHA data.

MASON is another dummy variable. This variable has a value of one if the frame of the major improvement at the residential location is masonry and zero if not. Its proper sign where factory fabricated frames are excluded has not been investigated by the writer.

MODUR is the term of the mortgage in years. All mortgages in this study appeared to be of the straight term type such that the mortgagor pays a fixed sum toward principal and interest each month. The cost per dollar of the mortgage to the mortgagor is viewed as the discounted value of his future interest payments less the discounted expected value of the asset he possesses upon full repayment of the mortgage loan at its maturity. Only if the mortgagor's subjective rate of interest is greater than the mortgage rate of interest will the mortgage actually be taken out. Thus, either an increase in the mortgage term or a decrease in the mortgage rate of interest would reduce the costs of purchasing bundles of housing characteristics relative to the costs of purchasing alternative characteristics bundles. The quantities demanded and offer prices for housing bundles would therefore increase. In the present study market interest rates were almost completely invariant for the sample within any one year and, for that matter, did not change greatly over the four year period being considered. Accordingly, only mortgage duration, whose sign is expected to be positive, is employed as an explanatory variable.

At this point, two peculiarities of FHA - insured mortgage loans relative to conventional mortgage loans should be recognized. First, the FHA requires that issuers of FHA - insured notes receive cash payments equal to the full face value of the note. That is, the cash payment cannot register any discount or premium which may be attached to a note when the note specifies a rate of interest below or above the prevailing market rate of interest. If, for example, a discount is in order, the mortgagor absorbs it because the seller inflates the price of the house by an amount reflecting the discount and then passes it on to the issuer of the note. Therefore to the extent that interest rates on FHA loans in Chicago from 1964 through 1967 were less than those on conventional loans, the observed price of a given bundle of housing characteristics in this study's data would be greater than the observed price of a similar housing bundle financed by a conventional loan. A cursory survey of prevailing rates during the period makes it appear that FHA rates after 1965 were about one-half of one percent less than those for conventional loans. Thus, for a given bundle of housing characteristics, the absolute magnitudes of the estimated coefficients of the explanatory variables in our results are somewhat greater than would be the case if the house purchase was financed by a conventional loan. In particular, the absolute magnitudes of the air pollution coefficients are somewhat greater.

The above tendency is reenforced by the fact that FHA-insured loans typically have only a small or nonexistent prepayment penalty, whereas conventional loans often have penalties amounting to as much as six months' interest. One would expect the positive value of the lack of an FHA prepayment penalty to be capitalized into the sale price of the property. However, it must be emphasized that neither this peculiarity nor that pointed out in the previous paragraph introduce a statistical bias to our estimates. Our results are fully applicable to residential property sold on FHA terms, but some slight downward revision must be made in the absolute magnitudes of the estimates obtained from the disaggregated FHA data if inferences are to be made about properties sold on conventional terms.

NHAGE is the average age of the residential properties in the 160 acre quarter - section in which the residence is located. Values for this variable were calculated from [53] where typical ages in each quarter section are presented. In cases where more than one typical age was presented, the weight each age received was the inverse of the number of "typical"ages. Given that bundles of housing characteristics generally provide less utility as they age, the demand for these bundles tends to decline. Price thus decreases and by the filtering process the bundles are thought to become available to households having relatively low incomes. This variable for the disaggregated FHA data is therefore a representation of the same phonomena as DILP6 is for the census data.

NWPP6 is the percentage of total population that was non-white in census tract as of April, 1960. As was explained for BLK66, the meaning of the coefficients obtained for this variable is not clear.

OLDER is the percentage of residential units in a census tract more than twenty years old in April, 1960, It is intended to measure the same phenomenon as HSAGE does in the disaggregated FHA data.

PMM refers to the minimum arithmetic mean monthly dosage during a given year of suspended particulates in micrograms per cubic meter per forty-eight hours. As is the case for all the air pollution variables, the two digits following the acronym refer to the year in which the air pollution dosages occurred. Thus, PMM64 is the minimum arithmetic monthly mean dosage of suspended particulates in 1964.

PMX is the maximum arithmetic mean monthly dosage during a given year of suspended particulates.

PRT is the annual arithmetic mean monthly dosage of suspended particulates during a given year.

PSK is the third statistical moment or skewness of the distribution of arithmetic mean monthly suspended particulate dosages during a given year divided by the annual arithmetic mean.

PVR is the second statistical moment or variance of the distribution of arithmetic mean monthly suspended particulate dosages during a given year divided by the annual arithmetic mean.

SMM is the minimum arithmetic mean monthly dosage during a given year of sulfur dioxide in milligrams per one hundred square centimeters per forty-eight hours.

SMX is the maximum arithmetic mean monthly dosage of sulfur dioxide during a given year.

SQIDX is an index of public elementary school quality in the residence's school district. The scale of the index has four factors with the digit 1 representing the highest quality and 4 the lowest. The source of the index and the explanation of its construction is to be found in [54, p. 146]. Generally speaking, the index is a composite of student performance, teacher quality, funding, and similar factors. If householders do attach positive utility to "good" schools, this variable's coefficients will assume negative signs.

STORY is the number of full stories in the major improvement at the residential location. This variable was employed only when MAINT was the dependent variable. The writer has given no attention to the sign, one might expect the variable to assume. In fact, it was included in the MAINT regressions only because it consistently proved to have a statistically significant coefficient.

SSK is the third statistical moment or skewness of the distribution of arithmetic mean monthly sulfur dioxide dosages during a given year divided by the annual arithmetic mean.

SVR is the second statistical moment or variance of the distribution of arithmetic mean sulfur dioxide dosages during a given year divided by the annual arithmetic mean.

SUL is the annual arithmetic mean monthly dosages of sulfur dioxide during a given year.

TCINC, defined as the current monthly family income of the residential property's owners, is the income variable always employed in the empirical analysis of the next section. Unfortunately, it is at best an imperfect proxy for the permanent income measure properly used as an argument in the offer function for bundles of housing characteristics. 20/ In broad terms, the use of current rather than permanent income can bias downward and make inconsistent the estimate of the positive effect upon offer price of additional income. Fundamentally, a measurement error is introduced because lagged income variables which would affect offer price are not included as regressors. If, as seems reasonable, negative transitory incomes appear more frequently in groups with low current incomes and if positive transitory incomes appear more frequently in groups with high current incomes, then the variance of measured income exceeds that of permanent income. Cross-sectional least squares estimates based upon current income would therefore exhibit a downward bias.

The income elasticity estimates appearing in the next section are generally slightly greater than .30. This is less than half the most widely accepted estimate of about .90 found, for example, in Lee [47] and Anderson and Crocker [1]. However, it nearly duplicates the .35 cross-sectional estimates for measured income elasticity reported by Reid [46, p. 9] and Malone [48], and approximates that of .41 found by Atkinson [49]. If FHA - insured mortgages are more likely to be obtained by better "credit risks" and if these individuals have lower measured income elasticities of demand for bundles of housing characteristics than does the population of residential property buyers, then such differences as remain between the current income elasticities obtained here and elsewhere could be accounted for.

Measures of what the FHA terms "effective income"--the FHA estimate of the household's earning capacity before deduction for Federal income taxes that is likely to prevail during the first third of the mortgage term -- are present in the data. When these measures of what at first appearance looks like permanent income were substituted for current income as regressors, an income elasticity of about .20 was obtained. This is to be expected since without exception the effective income measures when viewed as measures of permanent income implied that the transitory component of current income for 1,288 families was negative! None of the income elasticity results reported here are peculiar to the City of Chicago, for similar results were obtained with disaggregated FHA data in Chicago's north shore suburbs as well as in so-called downstate areas. Faced with these results, a search was then undertaken to find an instrumental variable, but no variable could be found which could reasonably be said to be correlated with permanent income, independent of the measurement error admittedly present in TCINC, and independent of other explanatory variables.

In and of itself, the measurement error involved in TCINC is of no particular relevance to this study. It is of concern only because such error can also bias the estimates of the effects of air pollution dosages upon offer price. If, for example, permanent income and damages are strongly and positively correlated and if the estimated effect of permanent income is biased downward, then the air pollution damage estimates will tend to be biased upward (closer to zero) because least squares procedures will cause the air pollution coefficients to register some of the positive impact of permanent income upon offer price. More or less economic theoretic arguments, e.g., [49, pp. 238-243], appear at several places in the pollution economics literature purporting to show that air pollution damages and (permanent) income are positively related. However, none of the previous work dealing with the covariation between air pollution dosages and residential property values perceived any collinearity between air pollution and permanent income requiring substantial adjustments in either estimates or estimation methods. 21/ Furthermore, current and permanent income are quite strongly and positively correlated. Thus if one income measure is collinear with air pollution dosages, it is not unreasonable to expect the other to be. However, employing the Farrar-Glauber [50] tests, no statistically significant collinearity was apparent between TCINC and the air pollution variables in the present study. We therefore proceed as if the admitted bias present in the estimated income coefficients does not affect the air pollution coefficients.

The Results

In order to facilitate evaluation of the results presented in this section, Table 3 immediately below presents the arithmetic means and standard deviations (in parentheses) of the original values of all the variables except the air pollution variables listed in the previous section. Table 4 presents the same thing for the various air pollution variables. In both tables, the columns refer to the year in which the property sale occurred. For certain of the variables, a column thus provides an idea of year to year variations in a variable or variables for a given set of residential locations, while the rows show how a given variable changes with changes in the set of residential locations. Table 5 presents simple correlation coefficients for the natural logarithms of some of the variables playing an important role in the material to follow.

The wide variety of economic-theoretic and empirical hypotheses raised in the previous sections makes organization of this section around a central theme rather difficult. Accordingly, an attempt is made to start from the

Table 3. Arithmetic Means and Standard Deviations of Nonpollution Variables

Year Variable	1964		1965		1966		1967	
Number of observations	245		212		345		386	
COSTS	15789	(2896)	16208	(3276)	16495	(3295)	17154	(3910)
PSITE	3256	(1104)	3402	(1237)	3310	(1031)	3272	(946)
FHAVL	15534	(2771)	16203	(3198)	16446	(3203)	17038	(3730)
MAINT	9.63	(2.34)	9.46	(2.54)	9.02	(2.31)	10.25	(3.22)
TAXES	25.84	(8.23)	26.11	(8.59)	26.42	(8.32)	25.71	(7.77)
MVAL6	16490	(3309)	16528	(2718)	16174	(3071)	15695	(3020)
MVL66	18379	(3586)	18139	(3905)	18056	(3531)	17274	(4230)
ALMNM	.03	(.18)	.02	(.14)	.06	(.23)	.15	(.35)
BLK66	37.21	(25.48)	28.06	(21.97)	28.78	(21.65)	34.24	(26.09)
BRICK	.67	(.47)	.70	(.46)	.71	(.45)	.60	(.49)
CRIMX	.80	(.73)	.77	(.54)	.83	(.58)	.90	(.62)
DEXPW	10.11	(9.22)	9.90	(8.70)	9.72	(8.57)	8.22	(6.05)
DLAKE	42.52	(22.23)	45.86	(18.67)	44.56	(18.82)	47.18	(16.22)
DILP6	.52	(1.13)	.57	(.97)	.54	(1.26)	1.67	(2.09)
DISSM	100.15	(20.92)	100.20	(19.51)	98.78	(19.70)	100.94	(20.04)
DTRAN	10.24	(10.19)	9.62	(7.69)	9.66	(7.63)	7.70	(4.24)
DTTYP	.90	(.30)	.95	(.21)	.95	(.22)	.95	(.22)
FRAME	.22	(.41)	.43	(.50)	.29	(.46)	.39	(.49)
HSAGE	28.27	(16.43)	30.85	(13.75)	32.40	(15.80)	31.56	(21.34)
INCID	238	(50)	251	(43)	251	(40)	252	(46)
INCM6	7622	(1153)	7714	(1202)	7697	(990)	7112	(1372)
INC66	8821	(1302)	9004	(1107)	8892	(1190)	8496	(1539)
INDUT	5.31	(6.21)	4.72	(5.69)	4.85	(6.50)	5.87	(6.90)
LIVAR	1181	(293)	1213	(334)	1194	(279)	1268	(312)
LOTSZ	NA		4620	(3142)	4347	(2871)	4154	(1168)
MASON	.76	(.43)	.56	(.50)	.69	(.46)	.60	(.49)
MODUR	287	(52)	289	(44)	293	(46)	292	(56)
NHAGE	44.08	(13.57)	46.34	(13.43)	47.31	(12.95)	49.68	(12.69)
NWPP6	12.39	(19.32)	10.86	(17.26)	10.35	(17.04)	20.05	(26.88)
OLDER	65.47	(51.79)	71.36	(57.45)	72.83	(59.85)	76.77	(59.61)
SQIDX	2.52	(.87)	2.27	(.73)	2.30	(.74)	2.39	(.71)
STORY	1.36	(.51)	1.36	(.51)	1.41	(.49)	1.40	(.50)
TCINC	788	(232)	806	(222)	864	(271)	918	(245)

Table 4 **Arithmetic Means and Standard Deviations of Representative Pollution Variables**

Year / Variable	1964		1965		1966		1967	
PRT64	131.43	(11.38)	128.05	(10.51)	130.22	(10.75)	130.09	(10.24)
PRT65					125.72	(9.59)	126.04	(9.94)
PRT66							138.54	(10.24)
SUL64	32.74	(17.54)	32.90	(17.47)	33.89	(19.76)	33.43	(20.22)
SUL65					37.38	(15.06)	37.30	(16.82)
SUL66							50.70	(11.59)
PMM64	90.67	(6.98)	91.20	(7.10)	92.01	(7.11)	93.38	(8.14)
PMM65					98.99	(8.63)	98.94	(8.01)
PMM66							99.09	(15.21)
SMM64	11.01	(4.61)	11.17	(4.58)	11.14	(4.60)	11.45	(5.44)
SMM65					16.19	(5.60)	15.83	(7.22)
SMM66							15.80	(3.36)
PMX64	155.37	(15.73)	156.30	(15.40)	158.30	(16.91)	157.32	(18.15)
PMX65					155.20	(8.35)	157.79	(12.20)
PMX66							160.32	(14.57)
SMX64	64.89	(36.46)	64.10	(35.24)	66.20	(39.85)	63.27	(40.28)
SMX65					63.84	(24.97)	63.98	(28.53)
SMX66							112.32	(24.84)

Table 5. Simple Correlation Coefficients for 1967 Observations.

	PSITE	BLK66	DISSM	DLAKE	HSAGE	INC66	INDUT	LOTSZ	MODUR	TCINC	PRT64	PRT65	PR
COSTS	.578	-.140	.398	.283	-.561	.357	-.128	.326	.716	.513	-.282	-.305	-.3
PSITE		-.130	.328	.080	-.234	.111	-.140	.700	.355	.297	-.179	-.115	-.1
BLK66			.010	-.029	.129	-.182	-.238	-.271	-.111	-.122	-.020	.001	-.0
DISSM				.231	-.515	-.001	-.147	.298	.420	-.014	-.461	-.371	-.29
DLAKE					-.304	.234	-.020	.109	.170	-.059	-.522	-.431	-.28
HSAGE						.168	.078	-.160	-.660	-.234	.119	.044	.0
INC66							-.167	.085	.089	.221	-.464	-.654	-.63
INDUT								-.098	-.039	-.128	.152	.198	.09
LOTSZ									-.058	.256	-.164	-.112	-.1
MODUR										.240	-.237	-.190	-.19
TCINC											-.129	.044	-.16
PRT64												.952	.96
PRT65													.96
PRT66													
SUL64													
SUL65													
SUL66													
PVR66													
PSK66													
SVR66													

Observations.

INDUT	LOTSZ	MODUR	TCINC	PRT64	PRT65	PRT66	SUL64	SUL65	SUL66	PVR66	PSK66	SVR66	SSK66
-.128	.326	.716	.513	-.282	-.305	-.314	-.249	-.242	-.369	.303	.376	.329	.233
-.140	.700	.355	.297	-.179	-.115	-.128	-.163	-201	-.262				
-.238	-.271	-.111	-.122	-.020	.001	-.003	-.034	.028	.163				
-.147	.298	.420	-.014	-.461	-.371	-.296	-.713	-.806	-.792				
-.020	.109	.170	-.059	-.522	-.431	-.284	-.480	-.376	-.530				
.078	-.160	-.660	-.234	.119	.044	.038	.348	.441	.511	-.654	.425	-.430	.028
-.167	.085	.089	.221	-.464	-.654	-.636	-.373	-.300	-.126				
	-.098	-.039	-.128	.152	.198	.091	.131	.131	.112				
		-.058	.256	-.164	-.112	-.114	-.225	-.238	-.226	.236	-.099	.099	.515
			.240	-.237	-.190	-.190	-.378	-.415	-.426	.438	-.331	.383	.043
				-.129	.044	-.163	-.139	-.120	-.142	.079	-.162	.070	.058
					.952	.960	.793	.720	.604				
						.967	.784	.712	.542				
							.708	.609	.457	.068	.523	-.485	-.081
								.963	.811				
									.913				
										-.480	.712	-.968	-.080
											-.296	.348	.045
												.663	.024
													.085

simpler and more familiar formulations, deal with the estimation problems associated with these formulations, and only then move on to the less familiar.

All regressions reported in this section are multiplicative. The coefficients of the explanatory variables are therefore elasticities showing the percentage change in the dependent variable associated with a one percent change in the explanatory variable. Throughout the assumption is made that the marginal cost of supplying housing characteristics is constant.

Table 6 presents the results when, except for the 1964 observations, the annual arithmetic mean air pollution dosages of the immediately preceeding year are regressed on the FHA-insured sales taking place during a given year. For 1964, the dosages and the sales coincide in time. The figures in parentheses are the associated coefficients' standard errors. $\overline{R}^2$ is the portion of unexplained variance adjusted for degrees of freedom, S is the equation's standard error, n is the number of observations, and ΔDAM is the marginal capitalized damages to the representative residential property due to air pollution dosages. The latter number is evaluated at the arithmetic means of the dependent variable and the air pollution variables. It represents the marginal capitalized damages for each additional ten micrograms per cubic meter per forty-eight hours of suspended particulates plus an additional one part per million per forty-eight hours of sulfur dioxide.

The sum of Table 6's air pollution coefficients has the expected sign in each of the four years. As is apparent from a glance at Table 5, the two pollution variables are highly correlated in each year. Application of the Farrar-Glauber [50] test to each of the pairs of pollution coefficients in Table 6 always revealed statistically significant multicollinearity. Further statistical evidence of serious multicollinearity in the pollution

Table 6. Initial Regressions. Dependent variable, ln (COSTS)

Year Variable	1967		1966		1965		1964	
Constant	5.3568	(.9143)	5.0227	(1.1631)	4.8396	(1.0461)	5.4809	(1.1126)
ln(PRT64)					-.2207	(.1616)	-.4343	(.1242)
ln(SUL64)					-.1304	(.0270)	.0552	(.0191)
ln(PRT65)			-.5025	(.1569)				
ln(SUL65)			.1298	(.1291)				
ln(PRT66)	-.3537	(.1051)						
ln(SUL66)	.0296	(.0593)						
ln(TCINC)	.3570	(.0336)	.3396	(.0636)	.2732	(.0366)	.3393	(.0291)
ln(LIVAR)	.2634	(.0385)	.2694	(.0265)	.1873	(.0388)	.2726	(.0372)
ln(HSABE)	-.0438	(.0080)	-.0490	(.0107)	-.0448	(.0157)	-.0474	(.0133)
ln(MODUR)	.3889	(.0448)	.3796	(.0479)	.3350	(.0627)	.2029	(.0533)
ln(LOTSZ)	.0483	(.0200)	.0744	(.0229)	.0489	(.0258)		
ln(SQIDX)	-.0648	(.0205)	-.0808	(.0208)	-.1169	(.0317)	-.0609	(.0229)
ln(NHAGE)	-.1106	(.0318)	-.1361	(.0635)	-.1252	(.0574)	-.1458	(.0531)
ln(DISSM)	.0471	(.0372)	.0453	(.0368)	.0543	(.0420)	-.0150	(.0595)
ln(DLAKE)	-.0236	(.0069)	-.0214	(.0093)	.0291	(.0234)	-.0186	(.0101)
FRAME	-.0761	(.0149)	-.0955	(.0164)	-.0818	(.0188)	-.0796	(.0203)
DITTYP	.1383	(.0288)	.1228	(.0312)	.1347	(.0405)	.1432	(.0261)
R2	.7711		.6819		.7034		.6022	
S	.1126		.1155		.1133		.1147	
n	388		345		212		244	
Δ DAM	$428.18		$601.45		$344.12		$510.79	

measures is presented in Table 7 below, where each of the pollution coefficients refers to regressions in which the other pollution measure was not included. For example, ln(PRT66) refers to a regression similar to that for 1967 in Table 6 except that ln(SUL66) was dropped. As is to be expected, standard errors, s, of the coefficients, $\hat{b}$, in Table 7 tend to decrease, $\bar{R}^2$ decreases very litte, and the signs of the SUL coefficients change from positive to negative.

Table 7. Initial Regressions with One Pollution Variable.

	1967		1966		1965	
	ln(PRT)	ln(SUL)	ln(PRT)	ln(SUL)	ln(PRT)	ln(SUL)
b	-.5337	-.1920	-.4163	1.2739	-.3172	-.1795
s	.0975	.0472	.1438	.1105	.1249	.0267
$\bar{R}^2$	.7683	.7555	.6792	.6781	.7006	.7003

If one has inspected the simple correlation coefficient of Table 5, it would appear that the existence of multicollinearity in the initial regressions of Table 6 is not limited to the air pollution measures. In particular, the high simple correlation coefficients between ln(DISSM), ln(DLAKE), and the pollution measures almost guarantee that dropping the former two variables from the initial regressions will affect the air pollution estimates. Table 8 shows what happens when this is done. As expected, the standard errors of the air pollution coefficients in Table 8 differ from those in Table 6. However, nowhere is the change so drastic as to justify acceptance of the null hypothesis that air pollution has no statistically significant effect on residential property values. Furthermore, the air pollution damage elasticities change relatively little so that the conclusions to be

Table 8. Initial Regressions Without ln(DISSM) and ln(DLAKE).

Year

Variable	1967		1966		1965	
Constant	5.4928	(.6914)	5.7158	(.8441)	4.5587	(.6042)
ln(PRt64)					-.3560	(.1097)
ln(SUL64)					.0306	(.0192)
ln(PRT65)			-.5587	(.1674)		
ln(SUL65)			.1489	(.1106)		
ln(PRT66)	-.3625	(.1057)				
ln(SUL66)	-.0828	(.0441)				
ln(TCINC)	.3258	(.0254)	.3105	(.0570)	.2919	(.0302)
ln(LIVAR)	.2492	(.0386)	.2259	(.0240)	.2318	(.0334)
ln(HSAGE)	-.0469	(.0061)	-.0475	(.0109)	-.0426	(.0072)
ln(MODUR)	.4008	(.0439)	.3611	(.0486)	.3434	(.0540)
ln(LOTSZ)	.0544	(.0239)	.0664	(.0232)	.0598	(.0220)
ln(SOIDX)	-.0640	(.0220)	-.0750	(.0205)	-.0716	(.0267)
ln(NHAGE)	-.1568	(.0457)	-.1683	(.0802)	-.1498	(.0619)
FRAME	-.0878	(.0148)	-.1004	(.0165)	-.0850	(.0190)
DTTYP	.1112	(.0299)	.1348	(.0320)	.1406	(.0 97)
$\bar{R}2$	.7664		.6731		.6853	
S	.1140		.1171		.1161	
n	388		345		212	
Δ DAM	$476.86		$667.30		$435.54	

drawn from Table 6 suffer no serious harm.

It is of course possible that the regressions of Table 8 are biased and inconsistent because variables which should be included, ln(DISSM) and ln(DLAKE), are not included. By partitioning these two variables, it is perhaps possible to avoid simultaneously the problem of the improper exclusion of a variable as well as the problem of multicollinearity. Accordingly, using 1966 sales, Table 9 presents regression coefficients for the air pollution variables where ln (DISSM) was partitioned on the basis of whether or not the residential location was more than ten miles from downtown Chicago. ln(DLAKE) was partitioned according to whether or not the location was more than three miles from Lake Michigan. The specifications of the equations in Table 9 are indentical to those of Tables, 6,7, and 8. If simple correlation coefficients can be taken as an indicator of the severity of multicollinearity, those of Table 10 would make it appear that the regression results of Table 9 are not severely stricken. This judgment is confirmed

Table 9. Regressions for DISSM and DLAKE Partitions.

Partition Variable	DISSM < 100		DISSM $\geq$ 100		DLAKE < 30		DLAKE $\geq$ 30	
ln (PRT65)	-.2781	(.0874)	-.5428	(.1974)	-.7247	(.1781)	-.4971	(.1579)
ln (SUL65)	.0816	(.1147)	.1752	(.0810)	.1594	(.0297)	.2117	(.0530)
ln (DISSM)	.0892	(.0665)	-.0164	(.0099)	.0186	(.0197)	.0716	(.0643)
ln (DLAKE)	-.0388	(.0194)	-.0422	(.0266)	-.0583	(.0178)	-.0042	(.0097)
$\bar{R}^2$	.6807		.6941		.6583		.7110	
S	.1159		.1147		.1342		.1085	
n	185		160		78		267	

Table 10. Simple Correlation Coefficients for DISSM and DLAKE Partitions

	DISSM < 100		DISSM ≥ 100		DLAKE < 30		DLAKE ≥ 30	
	ln(DISSM)	ln(DLAKE)	ln(DISSM)	ln(DLAKE)	ln(DISSM)	ln(DLAKE)	ln(DISSM)	ln(DLAKE)
ln(PRT)	-.2912	-.5136	-.0331	-.6464	-.3026	-.3290	-.4187	-.2556
ln(SUL)	-.5481	-.3534	-.4251	-.1313	-.6194	-.0915	-.4813	-.1679

when ln(DISSM) and ln(DLAKE) are dropped from the regressions of Table 9. The exclusion of these variables has no meaningful effect upon the standard errors of the air pollution coefficients.

On the basis of the immediately preceding manipulations, we conclude that the problem of multicollinearity between ln(DISSM) and the air pollution variables and between ln(DLAKE) and the air pollution variables does not justify acceptance of the null hypothesis that air pollution dosages have no impact upon residential property values. However, simple multicollinearity is not the only ground on which the initial results of Table 6 might be attacked. For example, ln(NHAGE) is supposed to be a composite measure registering the effects of neighborhood disamenities other than air pollution dosages. Yet it is quite possible that this variable does not register those disamenities which are correlated with air pollution dosages. If, for example, air pollution dosages tend to be higher where a substantial proportion of the surrounding land area is devoted to industrial uses and if these uses generate residential disamenities unrelated to air pollution, then the air pollution results of Table 6 may reflect the presence of these disamenities. Similarly, if air pollution dosages tend to be heaviest in areas populated by blacks or by low income people, and if others attach greater disutility to being near these groups than the utility the group members obtain being close to each other, then the air pollution variables would register the net negative effects upon the polluted area.

In Table 11, tests of these hypothesis are made by partitioning the sample. The first regression employs all 1966 property sales whose locations are at the center of a square mile having less than five percent of its area devoted to industrial uses. The last three regressions employ 1966 and 1967 property sales. The nature of the partitions is obvious from the previous list of variable definitions. Except perhaps for the BLK66 $\geq$ 75 regression, the air pollution coefficients of Table 11 are not at all inconsistent with those of previous tables. The results for BLK $\geq$ 75 can easily be due to the relatively small degrees of freedom and the fact that the standard deviation of ln()RT65) was onlv one percent of its mean while the standard deviation of ln(SUL65) was only six percent of its mean.

Multicollinearity and the improper exclusion of variables have not greatly bothered previous studies of this sort. Instead, as was earlier noted, concern has been exhibited about the introduction of aggregation bias due to the use of census data and measurement error caused by discrepency between the years of observed property sales and observed air pollution dosages.

Table 11. Regressions for INDUT, BLK66, and INC66 Partitions

Partition Variable	INDUT < 5		BLK66 < 25		BLK66 $\geq$ 75		BLK66 < 25 and INC66 $\geq$ 8495	
Constant	6.9438	(.9718)	8.0918	(.9804)	6.7449	(2.2473)	9.5894	(1.2727)
ln(PRT65)	-.5174	(.1810)	-.7331	(.1633)	-.5388	(.6278)	-.8263	(.2167)
ln(SUL65)	.1297	(0.385)	.2077	(.0457)	.2434	(.3618)	.2781	(.0704)
ln(TCINC)	.3313	(.0344)	.2835	(.0302)	.1600	(.0724)	.3537	(.0357)
ln(LIVAR)	.2177	(.0455)	.2015	(.0399)	.1066	(.0799)	.2401	(.0483)
ln(HSAGE)	-.0475	(.0155)	-.0521	(.0092)	-.0283	(.0227)	-.0561	(.0116)
ln(MØDVR)	.3673	(.0651)	.3783	(.0517)	.5955	(.1371)	.3388	(.0632)
ln(LØTSZ)	.0446	(.0288)	.0167	(.0268)	.0301	(.0899)	-.0091	(.0288)
ln(SQIDX)	-.0792	(.0243)	-.0291	(.0288)	.0162	(.0899)	-.0063	(.0406)
ln(NHAGE)	-.0227	(.0295)						
ln(DISSM)	.0689	(.0571)	.0446	(.0577)	.1151	(.2105)	.0520	(.0501)
ln(DLAKE)	-.0367	(.0199)	-.0346	(.0196)	-.1582	(.0984)	-.0462	(.0269)
FRAME	-.1039	(.0218)	-.0736	(.0178)	-.0440	(.0492)	-.0793	(.0218)
DTTYP	.1308	(.0386)	.1578	(.0299)	.0713	(.1054)	.1378	(.0320)
ln(CRIMX)			-.0456	(.0177)	-.1954	(.1770)	-.0459	(.0225)
$\bar{R}^2$	.6265		.7186		.7005		.7082	
S	.1190		.1199		.1288		.1159	
n	227		254		76		194	

Table 12 below compares the results for 1967 property sales when the years to which the air pollution dosages refer differ. The individual regressions therefore differ only in the year to which the air pollution dosages refer.

Table 12. Regressions Exhibiting Discrepencies in Air Pollution and Property Sale Data.

	1966		1965		1964	
Constant	4.1532	(1.0984)	4.1426	(1.1272)	4.1837	(.9626)
ln(PRT)	-.3998	(.1509)	-.4258	(.1223)	-.4557	(.1407)
ln(SUL)	.0601	(.0501)	.1223	(.0345)	.0808	(.0226)
$\bar{R}^2$	.7763		.7815		.7836	
S	.1116		.1103		.1098	
Δ DAM	$474.96		$601.45		$541.93	

The specifications of Table 12 differ slightly from those of table 6 in that ln(INCID) and ln(CRIMX) were included as explanatory variables while ln(NHAGE) was excluded.[23] Employing a method proposed by Tiao and Goldberger [55], an analysis of covariance was performed to test the null hypothesis of equality among years in the air pollution coefficients of Table 12. Table 13 presents the results of this test, where the columns refer to the combination of years for the air pollution dosages, F is the value of the F-test, and $\Sigma(T-K)$ is the sum of the regressions' observations less the number of variables. Only for ln(PRT64) and ln(PRT65) can the null hypothesis be accepted. However, on strictly intuitive grounds, it is difficult to discern any very important differences in any of the sums of the two pollution coefficients. Given our inability to describe with exactitude the synergistic properties of the two pollutants as the dosage of one or the other changes, it is certainly questionable whether an analysis of covariance can validly be carried out on the individual pollution coefficients.

Table 13. Analysis of Covariance of 1967 Disaggregated FHA Data

	ln(PRT65,66)	ln(PRT64,65)	ln(PRT64,65,66)	ln(SUL65,66)	ln(SUL64,65)	ln(SUL64,65,66)
$\Sigma(T-K)$	742	742	1113	742	742	1113
F	13.08	.78	12.80	43.17	33.70	26.21

A test similar to that above can be performed using census data, and, incidentally, insight can be acquired as to whether or not census results tend to exhibit aggregation bias. Table 14 presents regression results using census data. The air pollution dosages for each census tract correspond to the dosages found at an individual FHA-insured residential property near the center of the tract. Each column indicates the year to which the air pollution dosages refer.

An analysis of covariance of the air pollution coefficients of Table 14 yielded not a pair fulfilling the null hypothesis of equality of the coefficients for different mean annual pollution dosages. Again, however, the differences do not appear particularly important to the eye.

More interesting perhaps is a comparison of the air pollution coefficients obtained using the census data of Table 14 to the air pollution coefficients obtained using the disaggregated FHA data of Table 12. If the air pollution coefficients obtained using each of the two data sets are statistically equivalent, then the null hypothesis of an aggregation bias being present in the air pollution estimates obtained employing census data cannot be accepted. The results of an analysis of covariance to test this hypothesis are presented in Table 15. The columns refer to the variables common to both the FHA and the census regressions and the rows have the same meaning as in Table 13.

Table 14: Regression Results for Census Data. Dependent variable, ln(MVAL6).

	1966		1965		1964	
Constant	4.9049	(1.5794)	5.1712	(1.3420)	4.6524	(1.2799)
ln(PRT)	-.5515	(.2307)	-.5740	(.1841)	-.3754	(.1485)
ln(SUL)	.1839	(.1066)	.2410	(.0560)	.0654	(.0336)
ln(INCM6)	.6182	(1180)	.6036	(.0980)	.6124	(.1176)
ln(DILP6)	.0068	(.0050)	.0069	(.0049)	.0039	(.0049)
ln(OLDER)	-.0269	(.0133)	.0220	(.0119)	-.0240	(.0125)
ln(NWPP6)	.0103	(.0054)	.0112	(.0041)	.0127	(.0042)
ln(DISSM)	.1458	(.0670)	.1294	(.0583)	.1390	(.0588)
ln(DLAKE)	-.0425	(.0298)	-.0067	(.0274)	.0089	(.0301)
ln(SQIDX)	-.0769	(.0400)	-.0820	(.0408)	-.0751	(.0301)
$\bar{R}^2$	.5763		.5820		.5902	
S	.1391		.1352		.1339	
n	156		156		156	
Δ DAM	$418.73		$604.64		$426.19	

Table 15. Analysis of Covariance of Disaggregated FHA and Census Data.

	ln(PRT66)	ln(PRT65)	ln(PRT64)	ln(SUL66)	ln(SUL65)	ln(SUL64)
Σ(T-K)	514	514	514	514	514	514
F	1.39	8.69	1.97	6.60	16.60	2.17

In Table 15, ln(PRT66), ln(PRT64), and ln(SUL64) have values of F suitable for accepting the hypothesis of equality of the coefficients for these variables in the two data sets. The values of F for ln(PRT65), ln(SUL66), and ln(SUL65) do not fulfill the customary statistical criterion. Nevertheless, though the results of Table 15 do not entirely justify it, the writer is inclined to dismiss the possibility of the presence of aggregation bias in the air pollution coefficients obtained using census data. If any aggregation bias is in fact present in the census data, on the basis of these results one must attach a rather high value to the complete absence of bias in order to justify the effort required to compile and analyze disaggregated FHA or similar data.

The available data permitted a higher level of aggregation to community areas, a collection of several census tracts. Observations on forty-six community areas were used in regressions where MVL66 was the dependent variable and the natural logarithms of PRT65, SUL65, INC66, SQIDX, DISSM, DLAKE, and BLK66 were the explanatory variables. The sum of the air pollution coefficients was -.2527 and marginal capitalized damages were $438.54. However, the natural logs of the air pollution variables and ln(DISSM) were collinear. When the forty-six observations were partitioned by DISSM, both air pollution coefficients became statistically insignificant, though their sum continued to be negative. The question of whether or not meaningful air pollution damages estimates can be obtained when the unit of observation is larger than a census tract therefore remains open.

Census data and disaggregated FHA data are not the only data sources from which air pollution damages to residential properties could conceivably be estimated. For example a wide variety of local and state agencies assess property values and collect information on structural characteristics. In its first three columns, Table 16 presents results for 1967 FHA-insured property sales when ln(FHAVL) and ln(TAXES) serve as dependent variables. An analysis of covariance of the air pollution coefficients in the first column of Table 16 and the 1967 column of Table 6 reveals no statistically significant difference in the coefficients for ln(PRT66), though the coefficients for ln(SUL66) do exhibit such a difference. ln(FHAVL) would thus appear to be fairly satisfactory proxy for ln(C0STS).

Of the explanatory variables employed in the first column of Table 16, only ln(TCINC) or a similar measure of individual household income is unlikely to be obtainable from publicly accessible records. Though it has been shown in Anderson and Crocker [1, pp. 7-11] that a failure to include an income variable in an offer price equation results in a misspecification and its attendant biases and inconsistencies, it is of interest to obtain an empirical indication of the direction and magnitude of this bias. The second column of Table 16 therefore duplicates the specification of the first column except that no income variable has been included. Though increases in the pollution variables and decreases in the income variable tend to have a depressing effect upon offer price, they are inversely correlated in the data. Failure to include the income variable therefore increases the negative

Table 16. Regressions with ln (FHAVL), ln (TAXES), and ln (MAINT) as Dependent
Dependent

Independent	ln (FHAVL)		ln (FHAVL)		ln (TAXES)		ln (MAINT)	
Constant	4.7574	(.6285)	5.5620	(.5133)	-1.2340	(.6601)	1.8061	(.5247)
ln (PRT66)	-.3218	(.0927)	-.4185	(.1192)	-.2638	(.0974)	-.3081	(.1101
ln (SUL66)	-.0088	(.0472)	-.0090	(0506)	.1069	(.0496)	-.0179	(.0221)
ln (TCINC)	.3420	(.0219)			.3515	(.0230)		
ln (LIVAR)	.2115	(.0250)	.2094	(.0313)	.2011	(.0263)	.3910	(.0246)
ln (HSAGE)	-.0371	(.0052)	-.0407	(.0058)	-.0299	(.0055)	-.-305	(.0046)
ln (MODUR)	.4047	(.0392)	.4382	(.0411)	.0784	(.0412)		
ln (LOTSZ)	.0924	(.0217)	.0904	(.0234)	.1164	(.0228)		
ln (SOIDX)	-.0659	(.0176)	-.0856	(.0175)	-.0564	(.0184)		
ln (NHAGE)	-.0942	(.0351)	-.1213	(.0709)				
ln (DISSM)	.0285	(.0111)	.0574	(.0502)	.0256	(.0133)		
ln (DLAKE)	-.0219	(.0121)	-.0539	(.0152)	-.0263	(.0142)		
FRAME	-.0800	(.0131)	-.1201	(.0178)	-.0901	(.0137)	-.0818	(.0459)
DTTYP	.0997	(.0268)	.1496	(.0330)	.0418	(.0282)		
MASON							-.0438	(.0472)
BRICK							.0620	(.0219)
ALMNM							.0455	(.0203)
STORY							-.0461	(.0115)
$\bar{R}^2$	.8007		.7244		.6611		.5719	
S	.1001		.1121		.1051		.0975	
n	386		386		386		386	
Δ DAM	$399.12		$505.75		$ .44		----	

magnitude of the pollution coefficients. As is evident in the second column of Table 16, the bias that is introduced appears to be rather substantial.

The third column of Table 16 employs monthly payments of locally assessed property taxes as the dependent variable. Since the coefficients are elasticity measures, one would expect them to be identical to those found in the 1967 column of Table 6 if property taxes constitute the same proportion of the sale price of each and every property included in the sample. The naked eye says the coefficients differ. What the naked eye says is confirmed by an analysis of covariance. It would seem that the City of Chicago's residential property assessors understate the effects of air pollution.

In the fourth column of Table 16 is presented a regression in which ln(MAINT) is the dependent variable. These results are totally contrary to expectations since the air pollution coefficients have negative rather than positive signs attached to them. It certainly seems unlikely that increased air pollution results in reduced maintenance costs.

If rejection of the null hypothesis that air pollution is entirely neutral with respect to offer prices for residential properties is taken as the criterion of success in this study, then greater success can be had employing ln(PSITE) rather than ln(MAINT) or ln(TAXES) as a dependent variable. The results of doing this are to be seen in Table 17, where the columns represent the years in which the property sales occurred. In each year, the sum of the air pollution coefficients has the expected negative sign, and ln(PRT) is always statistically significant at the generally accepted levels. When one or the other of the air pollution variables is dropped from the regression, the remaining variable always assumes a negative

Table 17. Regressions for Site Values. Dependent Variable, ln (PSITE).

	1967		1966		1965	
Constant	9.8624	(1.8545)	7.0144	(2.0440)	-.2.1833	(2.9518)
ln (PRT64)					-.7128	(.2896)
ln (SUL64)					.1507	(.0551)
ln (PRT65)			-1.7651	(.31112)		
ln (SUL65)			.4516	(.0786)		
ln (PRT66)	-.7915	(.2691)				
ln (SUL66)	.0183	(.0920)				
ln (TCINC)	.2719	(.0380)	.2231	(.0387)	.2581	(.0416)
ln (LOTSZ)	.3633	(.0329)	.3750	(.0367)	.2565	(.0496)
ln (SOIDX)	-.1265	(.0392)	-.0514	(.0385)	-.2101	(.0607)
ln (INC66)	.2650	(.0883)	.2730	(.1176)	.4391	(.1367)
ln (DISSM)	.1937	(.1142)	.2156	(.1004)	.1879	(.0984)
ln (DLAKE)	-.0847	(.0270)	-.1102	(.0268)	-.0529	(.0470)
ln (CRIMX)	-.1088	(.0328)	-.0312	(.0233)	-.0012	(.0332)
ln (NHAGE)	-.0510	(.0457)	-.0343	(.0345)	-.0043	(.0603)
$\bar{R}^2$	.6194		.4983		.5138	
S	.1751		.1954		.2264	
n	386		345		188	
Δ DAM	$191.09		$323.85		$167.00	

sign and is always statistically significant. As was the case in the initial regressions of Table 6, ln (DLAKE) and ln (DISSM) tended to be somewhat correlated (correlation coefficients of .6 and .5 respectively with ln (SUL)) with the air pollution variables. The exclusion of ln (DLAKE) and ln (DISSM) had no appreciable effect on the standard errors of the pollution variables however; nor did the dropping of ln (PRT) and ln (SUL) have any appreciable effect upon the standard errors of the two distance variables.

The important feature of Table 17's air pollution coefficients emerges when the marginal capitalized damages falling from them are compared to the marginal capitalized damages falling from the air pollution coefficients of Table 6. In each of the three years for which regressions were run, the ratio of average marginal site damages to average site value was greater than the ratio of average marginal property damages to average property value. Similarly, the sum of the air pollution elasticities in the site value regressions is always greater than the sum of the same coefficients in the property value regressions. This outcome is consistent with our hypothesis that land or site values are more sensitive to air pollution dosages than are property values.

Results consistent with another earlier stated hypothesis are set forth in Table 18 where the 1966 property sale observations are partitioned by three distinct intervals of sulfur dioxide dosages. Thus, the observations included in the regression for the column headed SUL65 $\leq$ 25 include all those 1966 property sales from the disaggregated FHA data subject to an annual monthly average in 1965 of twenty-five parts per million or less of sulfur dioxide. The other two columns have analogous meanings.

Table 18. Regressions with ln (SUL65) Partitioned. Dependent Variable, ln (COSTS).

	SUL65 ≤ 25		25 <SUL65 < 45		SUL65 ≥ 45	
ln (PRT65)	-.6938	(.2219)	-.4938	(.1874)	-.3892	(.1062)
ln (SUL65)	.2001	(.0610)	.1135	(.1372)	.1521	(.0610)
Δ DAM	$802.83		$574.56		$427.31	
n	74		149		122	

Without bothering to resort to statistical tests to ascertain whether the observations for the three regressions of Table 18 were drawn from the same population, the above results appear to be consistent with our hypothesis of declining aggregate marginal damages. This assertion is statistically confirmed by the failure of an analysis of covariance to find a significant difference in the populations of the first and second columns. However, the covariance analysis rejected the hypothesis that the observations of the third column came from the same statistical population as the observations of the first two columns. Thus one cannot claim that all three sets of pollution coefficients correspond to different points on the same damage function. Nevertheless, the fact that this claim can be made for two sets of the coefficients provides some evidence that the aggregate marginal damage function declines at least over low and intermediate ranges of air pollution dosages.

Some further more or less impressionistic support for the hypothesis of a declining aggregate marginal damage function is presented in Table 19 where regression results for various partitions of the data set can be compared. The regression is listed in the dependent variable column. Reading across the rows, the columns refer respectively to the year in which the property sales occurred, the nature of the partition, the arithmetic mean of the natural logarithm of average annual monthly suspended particulates in the year preceding the earliest year of the regression's property sale observations, the arithmetic mean of the natural logarithm of average annual monthly sulfur dioxide in the year preceding the earliest year of the regression's property sale observations, the regression coefficient for ln (PRT), and, finally, the regression coefficient for ln (SUL). The horizontal double lines are intended to seperate partitions of different variables, while the starred regression coefficients indicate nonsignificance at the .05 level of the one-tailed t-test.

Since arithmetic means were not calculated for the original values of the variables in each of Table 19's partitions, the presentation can neither confirm nor deny the hypothesis of a declining aggregate marginal damage function. In addition, since no analyses of covariance were performed upon the coefficients of the regressions for the partitions, one cannot know whether observations were drawn from the same statistical population. Nevertheless, a cursory view of Table 19 makes it appear that damage elasticities tend to decline with increasing air pollution dosages. If the ratio of property value to air pollution dosages remains fairly constant throughout, a declining damage elasticity is consistent with a declining marginal damage function. The earlier results of Table 18 lead one to expect that, if calculated, the aggregate marginal damages of Table 19 would also tend to decline.

If it is in fact the case that marginal damages are greater at low than at high levels of pollution dosages, one would expect differential property values to be greater with differences in minimum dosages rather than with differences in maximum dosages. Table 20 shows six regressions for 1966 and 1967 property sale observations, where, except for the air pollution variables,

Table 19. Comparison of Logarithmic Means and Pollution Coefficients

Dependent Variable	T YEAR	Partition	$\overline{\ln (\text{PRT})}$	ln (SUL)	$\hat{b}_{PRT}$	$\hat{b}_{SUL}$
ln (COSTS)	67	MODUR ≥ 300	4.9206	3.8521	-.2376*	.0172*
ln (COSTS)	67	200 ≤ MODUR < 300	4.8746	3.5282	-.6913	.0866*
ln (COSTS)	66,67	BLK66 ≥ 75	4.9160	3.7983	-.5388*	.2434*
ln (COSTS)	66,67	BLK66 ≤ 25	4.8322	3.5091	-.7331	.2077
ln (COSTS)	66,67	BLK < .25 and INC66 ≥ 8495	4.8163	3.4053	-.8263	.2781
ln (PSITE)	66,67	BLK66 ≥ 75	4.9160	3.7983	1.2390*	-.0914*
ln (PSITE)	66,67	BLK66 ≤ 25	4.8322	3.5091	-2.1870	.3670
ln (PSITE)	66,67	BLK< 25 and INC66 ≥ 8495	4.8163	3.4053	-2.4939	.5108
ln (COSTS)	66	DISSM ≤ 100	4.8638	3.8110	-.2781	.0816*
ln (COSTS)	66	DISSM ≥ 100	4.7936	3.2438	-.5428	.1752
ln (COSTS)	66	DLAKE ≥ 30	4.8593	3.5174	-.4971	.2117
ln (COSTS)	66	DLAKE ≤ 30	4.7810	3.2516	-.7247	.1594
ln (COSTS)	65,66,67	HSAGE > 40	4.8640	3.4438	-.3795	.0217*
ln (COSTS)	65,66,67	20 ≤ HSAGE ≤ 40	4.8681	3.4107	-.2447	-.1506
ln (COSTS)	65,66,67	HSAGE < 20	4.8455	3.0848	-.7665	.1075*
ln (COSTS)	65,66,67	None	4.8593	3.3343	-.3560	.0227
ln (COSTS)	65,66,67	TCINC > 850	4.8639	3.3049	-.4475	.0374
ln (COSTS)	65	None	4.8553	3.3467	-.2207*	-.1304
ln (COSTS)	65	INDUT < 5	4.8464	3.3010	-.5174	.1297

the specifications were identical to the initial regressions of Table 6. The results presented in Table 20 are again fully consistent with the hypothesis of a declining aggregate marginal damage function. It should be recognized, however, that PMM and PMX are highly correlated with each other and with PRT, while SMX and SUL are also highly correlated. The regressions of Table 20 in which both minimum and maximum pollution dosages are included as regressors obviously exhibit this collinearity. Its presence inhibits any meaningful test of the hypotheses of the receptor's use of a minimax decision criterion in response to uncertainty about future pollution dosages.

In another of the hypotheses developed in a previous section about receptor responses to uncertainty, it was shown that if the receptor possesses a cubic utility function in air pollution, then the second and third moments of the annual distribution of pollution dosages become relevant to damages. Table 21 presents regression results for 1965, 1966, and 1967 property sale observations when various combinations of the first, second, and third moments of the distribution of air pollution dosages are included as regressors. Again, except for the air pollution variables, the specifications were indentical to the initial regressions of Table 6. All air pollution measures in Table 21 refer to the dosages occurring in the year immediately prior to the year of property sale observations listed in the columns.

The results presented in Table 21 are rather ambiguous. Nowhere does the inclusion of the second moment divided by the mean yield a statistically significant coefficient, though the signs conform to what was expected. This lack of significance for the measure of air pollution variance does not appear to be due entirely to multicollinearity problems for only ln(SVR66) and ln (SUL66) have an extremely high simple correlation coefficient. The

Table 20. Regressions for Minimum and Maximum Pollution Dosages

	1967		1967		1967		1966		1966		1966	
ln (PMM66)	-.2267	(.0527)			-.3139	(.0936)						
ln (SMM66)	-.1194	(.0350)			-.1283	(.0397)						
ln (PMX66)			-.2071	(.0719)	.0383	(.0954)						
ln (SMX66)			-.0209	(.0349)	.0607	(.0457)						
ln (PMM65)							-.4333	(.0919)			-.3758	(.1101)
ln (SMM65)							-.0557	(.0258)			-.0370	(.0399)
ln (PMX65)									-.4629	(.1301)	-.3642	(.1436)
ln (SMX65)									.0643	(.0223)	.1132	(.0375)
$\bar{R}^2$	.7722		.7641		.7721		.6916		.6669		.6915	
S	.1123		.1143		.1124		.1141		.1182		.1143	
n	388		388		388		345		345		345	
Δ' DAM	\$522.40		\$224.89		———		\$778.78		\$474.72		———	

Table 21. Regressions on Second and Third Moments of Pollution Dosages

	1967		1967		1966		1966		1965		1965	
ln (PRT)	-.4453	(.1133)	-.4570	(.1084)	-.5793	(.1700)	-.5495	(.1491)	-.3163	(.1693)	-.3206	(.1517)
ln (SUL)	-.1784	(.1525)	-0407	(.0459)	.1835	(.0515)	.1290	(.0295)	-.1521	(.0351)	-.0692	(.0276)
ln (PVR)	-.0195	(.0248)			-.0112	(.0304)			-.0265	(.0294)		
ln (SVR)	$-.14x10^{-4}$	$(.16x10^{-4})$			$.13x10^{-5}$	$(.49x10^{-5})$			$-.14x10^{-4}$	$(.21x10^{-4})$		
ln (PSR)	$-73x10^{-4}$	$(.36x10^{-4})$	$.68x10^{-4}$	$(.36x10^{-4})$	$-.64x10^{-5}$	$(.19x10^{-4})$	$-.87x10^{-5}$	$(.19x10^{-4})$	$-.72x10^{-4}$	$(.31x10^{-4})$	$-.75x10^{-4}$	$(.32x10^{-4})$
ln (SSR)	-.25x10-3	(.12x10-3)	$-.24x10^{-3}$	$(.12x10^{-3})$	$.74x10^{-4}$	$(.85x19^{-4})$	$.68x10^{-4}$	$(.79x10^{-4})$	$-.31x10^{-3}$	$(.14x10^{-3})$	$-.31x10^{-3}$	$(.13x10^{-3})$
$\bar{R}^2$	.7761		.7756		.6831		.6827		.7258		.7188	
S	.1121		.1122		.1149		.1150		.1124		.1125	
n	388		388		345		345		212		212	

drastic changes which always occur in the standard errors of ln (SUL) whenever ln (SVR) is added or dropped are strong evidence of multicollinearity however. 24/

The results for the skewness measures in Table 21 are somewhat more meaningful. In two of the three years, 1967 and 1965, each of the skewness variables is statistically significant. In the third year, 1966, where the skewness variables are not significant, the arithmetic means and standard deviations of the natural logarithms of the variables were no more than one-half the values of the other two years. In fact, the mean of ln (SSK66) was so small as to be practically monexistent. Further support for the meaningfulness of the skewness variables is offered by the fact that their signs, except for ln (PSK66), are in accord with expectations. For example, the sign of the arithmetic mean of ln (SSK66) is positive, implying that the air pollution distribution of 1966 for the locations of the 1967 property sale observations was skewed toward the higher pollution dosages. Given that disutility is attached to a greater density of pollution at the higher dosages, the sign of the coefficient for ln (SSK66) must be negative. Similarly, the sign of the arithmetic mean of ln (PSK66) is negative, implying a greater density for suspended particulates in 1966 at the lower dosages. As expected, the sign of ln (PSK66)'s coefficient is positive.

The presentation of the results of Table 21 finally completes the meaningful empirical findings of this study. Except for one, all the hypotheses, both theoretical and empirical, set forth in previous sections have been subjected to suitable estimation procedures and their results presented. The one hypothesis, the adaptive expectations hypothesis, for which no meaningful results have been or will be presented, was originally viewed as one of the study's major objectives.

The adaptive expectations hypothesis was tested by means of a nonlinear distributed lag program which was supposed to be capable of carrying out estimation where some variables in a cross-section were single-valued while others were vector-valued. 25/ In conjunction with property sales taking place in 1967, two different forms of the air pollution data were used. First, monthly observations of the mean daily dosages of each pollutant starting at the month of sale and extending as far back as January, 1964, for some runs were employed. Second, mean annual concentrations for 1966, 1965, and 1964 were employed. Whatever the air pollution data employed, monderately different results were obtained according to the starting values selected for the coefficient of the complete lag expression and the lag term of this expression. Never were statistically significant results obtained when monthly air pollution data were employed. However, results having moderate degrees of statistical significance were occassionally encountered when the mean annual air pollution data was used. In terms of statistical significance and intuitive reasonableness, the "best" results with these annual data are presented in Table 22. Except for the lag expressions introduced for the pollution variables, the equation specification was the same as in the initial regressions of Table 6. Seperate lag expressions were entered for suspended particulates and sulfur dioxide. The complete lag expression for each pollutant is $a_{i}\sum_{t=k=1}^{t=4} \lambda^{k} P_{k}$, where a is the regression coefficient, λ is the lag term, and P is the pollutant. The coefficients for all other explanatory variables closely approximated those of previous regressions. The percentage of variation explained was also similar.

Table 22. Example of Distributed Lag Results.

	Particulates	SO_2
a	-.0019 (.0010)	.0005 (.0004)
λ	.5094 (.0024)	.2966 (.8378)

At no time did λ PRT or λ SUL or λ PRT + λ SUL approach unity. Thus one can be fairly safe in asserting that some lag effect does appear to exist in the formation of expectations about future air pollution dosages. If one insists upon believing the results (the writer doesn't believe them) presented in Table 22, somewhere between fifty and eighty percent of the expectations held about air pollution dosages in any following twelve months are accounted for by the air pollution dosages of the previous twelve months. The fifty percent figure applies when one dismisses the insignificant λ for sulfur dioxide. It should be emphasized most strongly, however, that results having a degree of statistical resolution even approximately like those of Table 22 were obtained in only a minority of cases. In short, this study has added little if anything to knowledge about adaptive expectations in air pollution problems.

Footnotes

1. These same conclusions have recently been formally derived by Lind in [8] using an approach permitting weaker assumptions than in [7].

2. In particular, the negative influence of pollution is thought to be understated in ordinary least squares estimates. See [1, p. 13] for a statement and outline of a proof.

3. Though our concerns here are not normative, it is worth noting that receptor damages are not independent of the accuracy with which expectations about future air pollution dosages are formed. That is, realized receptor damages are not independent of the accuracy of previous receptor predictions. Realized damages are but present representations of positions taken in the past in response to expected air pollution dosages. Thus to the extent air pollution control instruments can reduce discrepencies between realized and expected dosages, these instruments can affect damage costs.

4. The expression "well-behaved" is a euphimism for some rather strong regularity conditions. In particular, $U(z_{i1})>U(z_{i2})$ if $z_{i1} > z_{i2}$ for all i and $z_{i2} = 0$ for at least one i. Furthermore, as the offer function is derived in [1], U must have strictly positive first order partial derivatives and continuous second order partial derivatives everywhere. Nevertheless, this representation is weaker than is usual, for, as Lancaster [14, pp. 156-157] shows, the formulation of U employed here need not be convex.

5. For more on this point, see Appendices A and B.

6. Since our offer function reads $c = f(z,y)$ this view of the identity of residential locations and consumer expenditures could mean that we have a greater number of different characteristics than we have price observations. However, as Lancaster [14,pp. 154-155] shows, given that U (z) is r - dimensional and convex, an n - dimensional slice of U(z) where ($n<r$) will have the same general properties as U(z) defined in r - dimensions. In any case, the problem of excessive characteristics does not seem applicable here. Kain and Quigley [17] could enumerate no more than twenty-one characteristics making a statistically significant contribution to house values.

7. This does not preclude ex post adjustments of behavior (adaptations of expectations) to realized errors in forecasts. The necessary conditions for the validity of certainty equivalence are ably presented in Malinvand [18].

8. The latter approach, termed the time - state preference approach to uncertainty, applies a discount rate to future contingent claims, where the rate discounts for futurity and for probability. A detailed presentation of the approach is available in Hirshleifer [19].

9. The following development leans very heavily upon the presentations in [20] and [21].

10. For an excellent short review of the principal features and implications of these alternative criteria, see Tisdell [23, pp. 14-41].

11. In accordance with the Hicksian composite goods theorem, we are assuming for simplicity that the implicit prices of the characteristics in the bundle always change proportionately. This will occur where the supply elasticities of these characteristics are equal and their elasticities of substitution for non-housing characteristics are the same.

12. It is interesting to compare the damage implications of the sufferer's use of a minimax criterion with the implications when an expected utility criterion is employed. Assuming for simplicity that only the first moment of the air pollution distribution is relevant, the expected utility criterion leads to a utility function of the form $E(U)= f[E(p)]$. Imagine a sufferer who with information $\{p,^* p_*\}$ uses the minimax criterion but with additional information might switch to the expected utility criterion. Now assume that if he had this additional information, the mean of the probability distribution of future dosages would prove to be $\bar{p}$, the midrange of the interval $\{p,^* p_*\}$. Given these assumptions, it is easily shown with the aid of diagrams similar to Figure 1 that if $\frac{\partial^2 c}{\partial z^2} > 0$ monotonically, then $E(p) > \bar{p}$; if $\frac{\partial^2 c}{\partial z^2} < 0$, then $E(p) < \bar{p}$; and if $\frac{\partial 2c}{\partial z2} = 0$, then $E(p) = \bar{p}$. The available evidence on the sign of $\frac{\partial^2 c}{\partial z^2}$ makes it appear non-negative. Data made available to the writer from the Chicago offices of the FHA indicate more or less constant marginal costs for provision of a wide variety of structural characteristics. The sign of $\frac{\partial^2 c}{\partial z^2}$ would thus be positive when it is recognized that the total costs of altering neighborhood characteristics probably increase exponentially because of the increasing marginal costs associated with the formation of coalitions involving increased numbers. In a recent study, deLeeuw [26] has found the elasticity of supply of "housing services" with respect to rents to be less than unity.

13. This is by no means meant to imply that all relevant issues with respect to the relation between air pollution dosages and utility have been thoroughly covered or even touched upon. For example, one might inquire about the implications for air pollution damage functions when expectations about future dosages differ among sufferers, when decision criteria differ among sufferers, and when both expectations and criteria differ. In short, air pollution damages under uncertainty are the joint result of expectations and decision criteria. Given that improved sufferer information about future air pollution dosages could be an important policy variable, the implications of differing expectations among sufferers about future dosages seems particularly interesting. First, uncertainty, and therefore expectations, is by its very definition of a subjective nature. Thus uncertainty can differ among sufferers even though they each possess the same information. Second, differing availability of information among sufferers implies differing degrees of uncertainty and thus considerable variability in expectations.

14. The developments in the following paragraphs owe much to readings of Lind [8], Muth [27], and Alonso [28].

15. The formal reasoning underlying these statements can be developed from Brandow [29] and Muth [30].

16. Records on the population of single family residential properties insured by the FHA in the City of Chicago from 1964 through 1967 were made available to this study. In the original records, the observations were grouped only by the time of their entry into these records. Assuming a more or less constant lag between time of sale and entry into the records, the observations were thus originally grouped by a sequence of dates. From this population an initial sampling selected every third entry. In a subsequent sampling every fifth remaining record was taken. The FHA records ultimately used for analysis thus constitute nearly forty-seven percent of the single family homes insured by the FHA in the City of Chicago from 1964 through 1967. A clear majority of these records refer to locations on the city's South Side.

17. Past and present administrators of the Chicago Air Pollution Control District have described the District's sampling and control program in a number of published papers and reports. A description is available, for example, in Stanley [37]. Additional information on the District's sampling and control program as well as the general nature of Chicago's air pollution problem is presented in Upham [39], Johnson [40], Stanley and Heller [41], and Northeastern Illinois Planning Commission [42].

18. The necessary meteorological data could not be obtained from the Chicago Air Pollution Control District.

19. The following provides a notion of the diversity of the data collected for each residential location: whether or not the improvement's exterior had a fiberboard finish, the inclusion of a dishwasher in the sale, the type of water service, wife's age, rate of commitments for felonies by neighborhood males from 1958 to 1962, the percentage of pupils in Sept. 1967 at the nearest elementary school below national reading norms, the floor space in the second closest shopping center, the zoning classification, and the seat miles of mass transit service available at the nearest rapid transit facility.

20. The economic-theoretic nature of the relation between permanent income and offer prices for bundles of housing characteristics is presented in Reid [45].

21. This is a bit of an overstatement since Ridker and Henning [2] carry out some residualization procedures because of perceived collinearity between their measures of permanent income and air pollution. Three comments are in order about their results. First, their air pollution data was grouped in that air pollution dosages over fairly wide intervals were assigned unique single values. This grouping would tend to increase the correlation between the air pollution and the income variables. Second, the air pollution damage coefficients Ridker and Henning obtained using a residualized permanent income measure as a regressor were within the standard error of the pollution coefficient when the income measure was not residualized. Third, as was noted in [1, p. 27], the Ridker-Henning

income coefficients, whether residualized or unresidualized, reflected the effect of mistakenly entering the median family income of a single census tract as zero.

22. It is worth noting here that an application of the t-test to any of the air pollution coefficients presented in this section means, given that $\hat{b}$ carries the correct sign, that $\hat{b}$ is significantly different from zero. We specify as the null-hypothesis the claim of some who argue that residential property values will not register any air pollution damages whatsoever. Thus, we are testing for coefficients permitting rejection of this null hypothesis. In short, we wish to follow a testing strategy which minimizes the chance of accepting the false hypothesis that an air pollution coefficient is significantly different from zero when it actually is not.

23. ln CRIMX and ln NHAGE appear to be collinear. ln INCID is not quite significant at the .05 level of the one-tailed t-test when ln NHAGE is included but it becomes significant (barely) when ln NHAGE is excluded.

24. One should note that the influence of variance in air pollution dosages upon expected utility could conceivably be non-linear in the variables and in the parameters. The earlier discussion of the cubic utility function could be made to suggest that the impact of the second moment upon expected utility varies with the level of air pollution dosages. Specifically, the behavior could change from risk-aversion to risk-preference as dosages fall, given that the marginal utility of less air pollution is always positive.

25. The program was written especially for this project by R.J. Anderson, Jr. of Purdue University's Department of Economics.

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Appendix A

Property Values and the Demand For Clean Air:
Cross Section Study for St. Louis

Kenneth F. Wieand

A Paper to be Presented at the
Committee on Urban Economics Research Conference,
September 11-12, 1970

Introduction

It is widely recognized that urban dwellers cannot pay for individual quantities of clean air and that they cannot, as individuals, charge polluters for damages caused by their effluents. Consequently, the control of pollution has passed into the hands of governmental agencies which are often empowered to set air quality standards and to enforce these standards by levying fines or instituting legal action against violators.

Officials in these agencies have discovered that the absence of market determined prices and quantities of air quality prevents them from determining the level of air quality where total benefits accruing from cleaner air neither exceed nor fall short of the required costs. Because this information is necessary to achieve an economically efficient use of resources, economists have attempted to measure indirectly the revealed preference for air quality.

One strategy for determining the demand for air quality argues that the negative effects of air pollution will be capitalized into the site value of residential housing. It follows that the demand for clean air can be determined by comparing differentials in property values in areas where pollution is high with property values where pollution levels are low. This paper summarizes an attempt to test the hypothesis that air pollution is negatively related to property values. The estimating equation is based upon a recent adaptation of the location model first proposed by Thünen [1]. Because of the lack of historical price data, the model is tested with data from a cross section of census tracts in the St. Louis area.

Ridker [2] and Crocker and Anderson [3] found a significant relationship between pollution and housing expenditures using the same data source as is used in the current study. This paper compares their studies to the one described here.

Pollution and the Price of Land

There is reason to believe that the property value-pollution level relationship will accurately reflect the deleterious effects of air pollution upon persons and property. Although pollution levels tend to be high in centers of production and population, factors such as wind direction and industry location distribute pollutants unevenly within surrounding housing markers. If housing consumers in a city recognize that pollution levels vary within the housing market, and if they prefer areas of higher air quality, sites in neighborhoods where pollution levels are low will be bid up relative to sites in neighborhoods where pollution concentrations are higher.

Because the concentration of air pollution is a characteristic of the housing unit's location, the price for clean air will be built into the locationally fixed factor, land. This can be illustrated by considering the investor who may build a house either in an area where pollution is high or in an area where it is low. Presumably, nonland costs are similar at the two sites. The investor will compare the discounted income streams of the two sites. He must subtract the added maintenance costs (or more rapid deterioration of the structure) from the income stream at the polluted site. The difference in the discounted income streams will be the price differential that the investor must pay for the site where pollution is low.

The appropriate model for testing the pollution-property value relationship can therefore be written $P_l = P_l\ (x_1, \cdots x_n,\ AP)$ where P_l is the per unit price of land, $x_1, \cdots x_n$ are determinants of land prices, and AP is an index of air pollution.

The Price of Land and the Price of Housing

As in most urban residential areas, the number of empty lots in the St. Louis housing market relative to all sites is negligible. Therefore, the price of land is not observable. Studies such as the one discussed here attempt to use the unit price of housing as a proxy for land prices. Housing prices are determined by the price of land and by the prices of nonland factors. Assuming that the housing market is competitive, and that households maximize profits at each site, the price of housing will vary directly with the price of land. To demonstrate this fact, Equation (1) relates housing expenditures to factor inputs.

$$(1)\quad P_h H = P_l\ L + P_n\ N$$

where P_h is the per unit price of housing, H is the quantity of housing, P_l is the price of land, L is the quantity of land, and P_n and N are the per unit price and the quantity of all nonland factors.

Assumptions of competition and profit maximization imply that a change in housing quantity will equal the sum of changes in land and nonland factors multiplied by their marginal productivities. Therefore, upon differentiating Equation (1) we derive

$$(2)\quad HdP_h = LdP_l + NdP_n$$

Equation (2) suggests that the price of housing will vary directly with the price of land. The price of housing is, therefore, an appropriate proxy for the price of land.

Total Expenditures, Unit Price, and Land Use Intensity

A major difference between the present study and the work of Ridker and of Anderson and Crocker is in the choice of the variable representing the price of housing. The latter two studies utilized the average rent and average market value statistics reported in Census of Population and Housing, 1960 as their dependent variables, while the author employed these same statistics and residential acreage to derive a measure of land use intensity.

Average rents and housing prices arise from transfers of housing services between economic units. In all transactions involving housing services, both those involving rentals over a period of time and those involving transfers of ownership, the total price is a multiple of the price per unit that the services can command, P_h, times the quantity of housing services that are involved in the transaction H.

In other words, average rents and average market values are not measures of the unit price of housing, P_h, but of total family expenditures on housing, P_hH. It can be shown that total family expenditures may move inversely with respect to the unit price of housing, if the quantity of housing consumed, H, changes in response to the change in P_h. This fact can be demonstrated by taking the derivative of P_hH with respect to P_h.

$$\frac{dP_hH}{dP_h} = \frac{HdP_h}{dp_h} + \frac{P_hdH}{dP_h} \qquad \text{or}$$

$$(3) \quad \frac{dP_hH}{dP_h} = H(1 + e)$$

where e = price elasticity of demand

If e is equal to -1, a change in P_h will leave total expenditures unchanged. If e is less than -1, total expenditures will vary inversely with changes in P_h. Recent research suggests that, in the case of individual units, e may be -1 or smaller. 1/ Therefore, $P_h H$ is not an adequate surrogate for P_h.

The use by past writers of family expenditures as a surrogate for unit price may result from an implicit assumption that the quantity of housing can be held constant between units by the addition to the list of regressors of variables which measure the quantity of housing. The quantity, H, of housing services depends upon floor space, layout and room design, materials used in construction of the unit, site size, and a number of other factors. A reasonably complete list of the components of H has been compiled by Kain and Quingley [4] If these data were available, they could perhaps be added to the list of regressors to hold variation of H constant. The 1960 Census did not make available the necessary detail, however, and H is not held constant.

Given an additional piece of information, the problems en-countered in using household expenditures can be avoided. Muth [5] has shown that the intensity of land use will vary directly with the price of land, hence with the price of housing, P_h. To briefly reconstruct his proof, return to Equation (1) and divide both sides by L.

$$(4) \quad \frac{P_h H}{L} = P_l + \frac{(N)}{L} P_n$$

Taking the derivative of Equation (4) in natural logs, one can prove that

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1/ Richard F. Muth, Cities and Housing, (Chicago: University of Chicago Press, 1969), p. 69.

$$(5)\quad d\ln\frac{(P_h H)}{L} = \left(1 + \frac{f_n}{f_1}\delta\right) d\ln P_h - \frac{f_n}{f_1}\delta\, d\ln P_n$$

f_1 = factor share of land

f_n = share of nonland factors

δ = elasticity of substitution

$\frac{P_h H}{L}$ will vary directly with the price of land, provided that the cost of nonland factors is constant throughout the labor market, because δ, f_n, and f_1 are positive. If δ is approximately 0.75 and the share of land is about 5 percent, relative changes in $\frac{P_h H}{L}$ will be about fifteen times the size of relative change in P_h, and about the same magnitude as changes in P_1. 2/

The primary measure of land use intensity employed in this paper was derived by multiplying average rentals plus 0.01 of owner estimated market values in each census tract by the number of each type of unit in the tract. (In the real estate profession 0.01 is the conversion rate used to appraise the rental of homes given their market value.) This figure measures the total monthly expenditure on housing in each census tract. The land area of each tract was computed, and a land use map was used to net out industrial and commercial acreage. Total monthly expenditure on housing was then divided by the residential land area measured in acres to arrive at monthly rent per acre. This variable, measured in natural logs, is named VALAND. VALAND is similar to the variable

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2/ *Ibid.*, p. 82, p. 55.

used by Muth in a study for south Chicago [6]. The lists of regressors used in the two studies are also quite alike, except for different measures of the age of housing and for the addition of the pollution variables in the present study. Predictably the two sets of results are quite similar.

A measure of average expenditures on housing (VALHOU), similar to that used by Ridker and by Crocker and Anderson, was also utilized as a dependent variable in an attempt to compare the performance of the independent variables in the three studies.

Determinants of the Intensity of Land Use

The independent variables used to try to explain variance in VALAND were the percent of standard units in a census tract (STDPER), average age of housing in the census tract (AVGAGE), percent of units built before 1920 in the tract (BEF20), percent white population in the tract (WHITE), average income in the tract (INCOME), distance in miles from the central business district (DISCBD), a dummy for proximity to major highways (HIWAY), a dummy for tracts in East St. Louis, Ill. (ESL), a dummy representing tracts on the boundary of a nonwhite area (BOUND), a dummy for proximity to industry (INDUS), and a dummy for proximity to commercial areas (COMM).

The first three variables were intended to explain the variance in VALAND resulting from changes in the unit price and quantity of the nonland portion of housing services. STDPER was included to isolate tracts where housing units have large amount of deferred maintenance; theory suggests that the coefficient of this variable is positive. AVGAGE measured the change in land use intensity over time. Decreasing transportation costs over time should make the coefficient of AVGAGE positive, but it could be negative if the variable picked up housing deterioration in older units. BEF20 measured technical change in newer housing such as expanded electrical circuitry, and gas and electric heating systems; the coefficient of BEF20 should be negative.

The next two variables measured neighborhood effects. The unit price of housing services was expected to rise in high income areas because of neighborhood amenities. It has been argued both that non-whites pay more and that they pay less than whites for equal quantities of housing. The coefficients of WHITE and BOUND could not be prejudged.

The final five variables defined the locational characteristics of the housing in a tract. Increased commuting costs tend to cause housing prices to decline as distance from the CBD increases. Physical seperation of East St. Louis should make ESL negative with respect to VALAND, whereas decreased commuting time near expressways should cause the coefficient of HIWAY to be positive. Proximity to industry may either increase or decrease VALAND: while commuting to work is easier, adverse external effects tend to drive down housing prices. Finally, units near commercial centers should command a higher price.

Data on sulphur pollutants and particulates are available on a spatial basis for a number of large cities. Ridker and Crocker and Anderson included sulfation, measured as sulphur trioxide, and suspended particulates, measured with hi-vol samplers, in their models. The paper under discussion utilized this information along with measures of dustfall and sulphur dioxide.

Like the other variables in the model, pollution levels vary spatially. Unlike the other variables, pollution concentrations also fluctuate over time. One can prove that the average concentration of a pollutant for a given period of time will accurately reflect the disutility of the pollutant's temporal distribution only if the marginal disutility of pollution is constant. Because the effects of pollution become more noticeable as concentrations increase, the assumption of constant marginal disutility does not seem warranted. Therefore, measures of the variance of the temporal distribution were also included. As the inclusion of these measures did not improve the performance of the pollution variables, the results obtained by using these variables are not presented here.

In the following presentation of results PAR stands for variables measuring suspended particulates, and SO_3 represents measures of sulphur trioxide. Existing data do not define either particulates or sulphur for each census tract in the study area. It was therefore necessary to estimate pollution levels for most tracts. Three methods of estimation were employed. sPAR and sSO_3 are variables obtained from isopleths (lines of equal actual value) found in the Interstate Air Pollution Study [7]. If the majority of a census tract lay between two isopleths, air pollution in the tract was assumed to equal the level prevailing on the boundary of the isopleth furthest from the CBD. These variables are similar to the proxies used by Ridker and by Crocker and Anderson.

Pollution measured in this fashion resembles a step function decreasing with distance from the CBD. One would expect pollution to decline continuously rather than in discrete jumps. Therefore, variables

iSO_3 and iPAR were estimated by interpolating in a straight line between isopleths. Finally, SO_3 and PARm were derived using data from isopleths, monitoring stations, and information on wind direction. SO_3 and PARm are expected to be the most accurate measures of the average levels of particulates and sulphur pollution for each census tract.

Empirical Results and Conclusions

Table I. reports the results obtained by regressing VALAND against regressors other than the pollution variables. All but HIWAY have signs consistent with a prior: expectation, all but BOUND, HIWAY, and COMM are significantly related to VALAND, and the magnitudes of the coefficients are reasonable.

Table II. lists the coefficients of the pollution variables. INDUS, COMM, and BEF20 are correlated with the pollution variables. For purposes of exposition, these three variables are omitted from the estimating equation.

The coefficients of pollution are negative in most cases but are not significantly related to VALAND. If either ESL or STDPER is omitted from the estimating equation, the coefficients of the pollution variables fall and become significant at the .05 level. The author doubts that multicollinearity between pollution and ESL and STDPER is responsible for the lack of a significant relationship between pollution and VALAND for two reasons. First, the standard errors of the pollution variables do not fall when ESL and STDPER are delected from the equation. Also, a subsample of 86 observations was selected by dropping all tracts in

East St. Louis and all tracts where the percent of standard units was small to determine whether ESL and STDPER obscured a significant relationship between pollution and VALAND. When the equation is tested using the subsample, the pollution variables behave as they do in the fully specified model.

Table III. lists the percentage changes in VALHOU and VALAND implied by 10 percent changes in some of the independent variables. Examination of Table III. shows that the relationships between VALHOU and the exogenous variables may be quite different from the relationships between these variables and VALAND. The coefficients of INCOME in the VALHOU regressions imply that the average value of housing in the study area is twice the size of the average yearly income, reflecting the fact that average housing prices in the SMSA were about $12,000 in 1960, and that average income was about $6,000. This relationship obviously reflects the fact that the elasticity of demand for housing is greater than zero. If relative changes in VALAND are fifteen times the changes in the price of housing, the 7.5 percent change in VALAND implies a 0.5 percent change in the price of housing. Consider also the changes in VALHOU implied by a 10 percent change in DISCBD. Three of the VALHOU regressions state that expenditures increase with distance when we would expect price to decline. Positive distance coefficients are consistent with a price elasticity of less than -1. The VALAND regression implies, on the other hand, that the price of land declines at 12 percent per mile and therefore, that the price of housing falls by a little less than 1 percent per mile.

This paper concludes that, although there is reason to believe that the price of housing services is negatively correlated with levels of air pollution, average rentals, market prices, or other measures of household expenditures are not adequate proxies for the price of housing. The output land ratio used in the study summarized here is a suitable proxy for the price of housing. When the output land ratio is regressed against levels of air pollution and other determinants of housing prices, the regression coefficients of the pollution variables are generally negative, but are not significantly different from zero at the .05 level.

The results of the study discussed above do not prove that the effects of air pollution are not harmful to urban dwellers. They do caution against using past findings to justify reductions in pollution levels.

Table I

Results of Regressing VALAND Against Variables Other than Pollution Variables

Variable	Coefficient (Standard Error)	Equation	
		4.2	4.3
STDPER		0.59 (.20)	0.59 (.20)
WHITE		-0.22 (.10)	-0.29 (.096)
AVGAGE		0.021 (.005)	0.021 (.005)
BEF20		-0.61 (.11)	-0.61 (.11)
INCOME		$0.6\mathrm{x}10^{-4}$ $(.2\mathrm{x}10^{-4})$	$0.7\mathrm{x}10^{-4}$ $(.2\mathrm{x}10^{-4})$
DISCBD		-0.116 (.023)	-0.122 (.023)
INDUS		-0.24 (.064)	-0.26 (.061)
ESL		-0.75 (.11)	-0.74 (.11)
HIWAY		-0.10 (.064)	
BOUND		0.12 (.080)	
COMM		0.091 (.077)	
R^2		0.64	0.62
Constant term		6.06	6.05
Degrees of freedom		126	129

Table II

Performance of Three Alternate Measures of Particulates and of Sulfur Trioxide in Explaining VALAND[a]

Variable	Coefficient	Standard Error	R^2
Sulfur Trioxide[b]			
S03	+0.0036	0.001	0.45
iS03	-0.0790	0.105	0.45
sS03	-0.0870	0.150	0.45
Particulates[c]			
PARm	-0.0015	0.0013	0.45
iPAR	-0.0011	0.0017	0.45
sPAR	-0.0012	0.0021	0.45

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[a] The specification used is as follows: VALAND = f(pollution variable, DISCBD, STDPER, AVGAGE, INCOME, WHITE, ESL)

[b] Measured in milligrams/hundred cubic centimeters

[c] Measured in parts per million

Table III

CHANGE IN VALHOU AND VALAND RESULTING FROM A TEN PERCENT CHANGE IN SELECTED INDEPENDENT VARIABLES: COMPARISONS OF WIEAND RIDKER, AND CROCKER - ANDERSON STUDIES

Independent Variable	Dependent Variable				
	VALHOU				VALAND
	Wieand owned and rented	Crocker - Anderson owned	Crocker - Anderson rented	Ridker owned	Wieand owned and rented
sSO3	- 1.0	- 2.0	- 0.2**	- 0.8	- 0.9**
SPAR		- 1.7	- 1.1		- 1.2**
INCOME	26.0	19.0	14.0	20.0	7.5
DISCBD	2.7**	- 4.0	1.0**	-1.3	-12.0
Built before 1940	- 1.0**	- 0.4	- 1.0	- 0.1	
Condition					
Percent Standard	5.1				6.6
Percent Dilapidated		16.0	13.2		
WHITE	- 2.8	- 1.2	- 0.8		- 2.3
AVGAGE					8.0
Median Rooms				9.0	
Constant ($\log_e$)	7.8	3.5	- 1.2	7.3	6.1
R^2	0.75	0.76	0.77	0.94	0.66
Degrees of freedom	139.0	236.0	174.0	167.0	139.0

**Not significantly different from zero at the 0.05 level.

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Appendix B

A COMMENT ON "PROPERTY VALUES AND AIR POLLUTION: A CROSS SECTION ANALYSIS OF THE ST. LOUIS URBAN AREA", BY KENNETH WIEAND

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I.

In accordance with principles of commentary reportedly established by Mark Anthony, we come not to praise what is praiseworthy in Wieand's paper, but rather to bury that which we believe to be in error. Ironically, we contend that Wieand's dismissal on logical grounds of our results and those of other investigators is fundamentally illogical.

In section II, Wieand's contention that we and others have used a theoretically incorrect dependent variable in our empirical work is evaluated. Then, in section III, some results from past property value studies not generally available are reported. In section IV, we present some new empirical evidence and offer some concluding comments.

II.

It is well-known that no universally accepted concept of a commodity exists in economics. Witness the economist's continuing concern with aggregation problems. To modify only slightly Bertrand Russell's definition of matter, a commodity is an abstraction defined in accordance with the purposes of a theory. The fundamental difference between Wieand's work and that of previous investigators lies in the adopted definition of commodity units. Wieand treats "housing" as a single commodity, whereas other investigators treat physically different properties as distinctly different commodities.

Following Muth (1969), Wieand defines observed "property values" or rentals to be identically the product of two unobservable magnitudes, the "price of housing services" and the "quantity of housing services". Under general assumptions, Wieand demonstrates that *if such a definition is adopted*, then it is the case that "property value" is neither equal to nor does it vary with the "price of housing services". This conclusion is unimpeachable, as is the logic to which it gives rise in the search for an observable proxy variable for the "price of housing services."

It does not follow, however, that property value is not a measure of price *under all definitions of* commodity units and attendant prices. Yet Wieand's criticism of our work rests, insofar as we are able to determine, on precisely this *non-sequitur*. To reject theorems based on one set of definitions and axioms because they happen to differ from one's own theorems based on a different set of definitions and axioms is at best unfair, and at worst bad logic.

The basis for the definitions of commodity units adopted in the work of previous investigators is Lancaster's (1966) formulation of consumer behavior and Baumol and Quandt's (1966) empirical work on the demand for abstract modes of transportation. We shall follow Lancaster.

The point of departure of these approaches to demand theory is that utility functions are defined not over traditionally conceived goods but rather over characteristics embodied in goods. The endowment of characteristics embodied in any good is presumed to be objective in the sense that all consumers perceive identical endowments in the same good. The allocation problem of the consumer is to consume that basket of goods which yields a mix and magnitude of characteristics maximizing his utility subject to his budget constraint. Assuming that goods and characteristics are linearly related (that the consumption technology is linear in Lancaster's phrase), the problem of the consumer may be formulated instructively as

$$\begin{aligned} &\text{Maximize:} \quad U(z) \\ &\text{subject to:} \quad p'x = y \\ &\qquad\qquad\quad z = Bx \\ &\qquad\qquad\quad z, x > 0 \end{aligned}$$

where z is an mxl vector of characteristics, p is an nxl vector of good prices, y is a scalar representing disposable money income, and B is an mxn matrix, the (i,j) element of which gives units of the ith characteristic obtained per unit of the jth good.

Consider the first-order conditions of consumer equilibrium,[1]

$$\begin{aligned} &U'\nabla_i z - \lambda p_i = 0 \qquad i = 1, \dots, n \\ (1)\quad & \\ &y - p'x = 0 \end{aligned}$$

where U' is the transpose of $\partial U/\partial z$, and where $\nabla_i z = \partial z/\partial x_i$. In the particular case of a linear consumption technology, $\nabla_i z = b_i$ is the ith column of the consumption technology matrix B.

If the conditions of the implicit function theorem are fulfilled by the implicitly defined vector function (1), there exists a unique explicitly defined vector function

$$(1a) \quad \begin{pmatrix} p \\ \lambda \end{pmatrix} = G(x, y)$$

The coordinate functions of G( ) which pertain to the price vectors of p may be interpreted as offer functions, giving prices offered for various quantities and income. These functions may be easily obtained directly from the first-order conditions, which by rearrangement and substitution yield

$$p_i = \frac{y\, U'\nabla_i z}{\sum_{j=1}^{n} U'\nabla_j z x_j} \tag{1b.1}$$

$$\lambda = \frac{\sum_{j=1}^{n} U'\nabla_j z x_j}{y} \tag{1b.2}$$

It is evident that the consumption technology is instrumental in the determination of marginal rates of substitution among commodities and hence is instrumental in determining demands. Note that a good has no identity of analytical significance apart from the characteristics which it yields. That is, goods which yield the same mix and magnitude of characteristics (goods for which the $\nabla_j z$ are identical) are perfect substitutes, on a unit for unit basis, and therefore bear the same price.

Equation (1b.1) is the theoretical underpinning for the statistical work with property values performed by us and others. Treating the observed average characteristics of each census tract as an observation on $\nabla_i z = b_i$, and median property value as an observation on p_i, we have attempted to explain differences in p_i in terms of y and b_i.[2] We thus treat each census tract observation as though it were an observation on a different commodity, while Wieand treats his observations as though they pertain to the same commodity.

III.

Some results from past studies of air pollution and residential property values are reported below in Table I. All investigators believe they have found statistically significant negative relationships between air pollution and property values. While the specifications differ in some respects, the air pollution results are remarkably uniform. That is, several replications have yielded very similar results, even though the structure of residential property markets among cities may differ widely. This is cause to be cautiously confident that some misspecification is not at base responsible for the observed negative relationship. However, since the pollution measures employed vary among these investigators, precise comparisons are not possible. Nevertheless, estimates of marginal property value losses vary between roughly $250 and $1000 per residential unit for ten to fifteen percent increments in the mean value of the various pollution measures used. Clearly these results and the conclusions which might be drawn from them differ from those of Wieand.

TABLE 1

Some Results of Past Studies of Air Pollution and Residential Property Values

Dependent Independent	Zerbe(1969)$^{++}$ In MPV	Zerbe(1969)$^{++}$ In MPV	Peckham(1970)$^{+}$ In MPV	Anderson-Crocker(1970)$^{+}$ In MPV	Anderson-Crocker(1970)$^{+}$ IN MPV	Anderson-Crocker(1970)$^{+}$ In MPV
CONSTANT	8.97	8.72	4.7821	-1.0705	-1.2325	1.1617
In PPT	--		-0.1155	-0.0627	-0.0641	-0.1698
In AMS	-0.1206	-0.0810	-0.0956	-0.0412*	-0.0649	0.0010*
In ITL	0.0041					
In MRM	0.2570	0.1260	0.44973	1.0996	0.8246	0.9064
In PPU	0.1720					
In RMFI	0.0001					
In RHPM	-0.0001					
In UKR	0.0062	0.00004*				
In JEW	0.0043					
In MOO	0.0029	0.0101				
In MFI		0.00004*	0.4867	0.7596	0.8130	0.9970
In DIS			0.0575	-0.0085*	0.0349*	-0.0312
In OLD			-0.0415	-0.0036*	-0.0464	-0.0213
In NWT			-0.0054*	-0.0004*	-0.0133	0.0321
In DLP			-0.0257	-0.0267	-0.0028*	0.0113*
R^2	0.942	0.923	0.766	0.877	0.916	0.790
S^2	0.0846	0.0488		0.0136	0.0114	0.0179
F	107.83	72.55	80.07			

*Not significantly different from zero at .05, 1 tailed test

+Regression included only observations on tracts where at least 75% of dwelling units were single family residential (generally suburban areas).

++Zerbe's equations were run in semi-logarithmic form.

NOTES TO TABLE I

PPT	annual arithmetic mean suspended particulates
AMS	annual arthimetic mean sulfation
ITL	percentage Italian
MRM	median number of rooms
PPU	people per dwelling unit, average
RMFI	residualized median family income
RHPM	residualized housing units per mile
OKR	percentage Ukranian
JEW	percentage Jewish
MOO	neighborhood occupational homogeneity
MFI	median family income
DIS	distance to central business district
OLD	percentage of housing units more than 20 years old
NWT	percent nonwhite
DLP	percent housing units dilapidated

Regressions reported in columns of this table are based respectively on data from Toronto (Ontario, Canada), Hamilton (Ontario, Canada), Philadelphia, St. Louis, Kansas City, and Washington, D.C.

IV.

If we may take it as established that the logical problem Wieand claims to have demonstrated in previous work is really no problem at all, it remains to determine the source of the difference between his results and those of other investigators. Although it is possible to entertain a near infinitude of dark suspicions of misspecification and other errors in giving empirical content to both theories, we believe that the divergence of results is a matter of chance pure and simple. This conjecture is based primarily on some regressions we have recently run using better data than the Census material on which all earlier work has been based.

The results in Table 2 are based on FHA data for individual property transactions in the City of Chicago.[3] Air pollution measures, which are interpolations from basic point measurements made by the Chicago Air Pollution Control District, refer to the mean annual suspended particulates and sulfur dioxide in the calendar year immediately preceding the year of property sale. As with our earlier work with census tract data, we have included measures of both suspended particulates and sulfur oxides in all equations even though the two tend to be so collinear that it is difficult to gauge the independent effect of each. In column 1 of the table, a specification analogous to that of Anderson and Crocker (1970) is employed, using as the dependent variable the actual price at which the transaction occurred. Column 2 reports the results for a regression in which the natural log of the FHA estimated price of the site for each transaction in the sample is used as the dependent variable. The results reported in column 3, where the natural log of the ratio of sale price to lot size

TABLE 2

Some New Results for Chicago - 1967

	(1) ln SALEP[a]	(2) ln SITEP	(3) $\ln \frac{SALEP}{LOTSZ}$
Constant	4.1532	7.0780	2.6615
PRT66	-0.3998	-0.0005	-0.0053
SUL66	0.0610*	0.0001*	0.0012*
TCINC	0.3305	0.0004	-0.00005*
LIVAR	0.2356		-0.00004*
HSAGE	-0.0465		-0.0039
MODUR	0.4063		0.0011
LOTSZ	0.0917	0.0002	
SOIDX	-0.0729	0.0543	0.0138*
INCID	0.0361		
DISSM	0.0321*	0.0003*	0.0028
DLAKE	-0.0202	-0.0024	-0.0011*
INC66	0.1216	0.00005	
FRAME	-0.0810		-0.1070
DTTYP	0.1004		-0.2906
NHAGE		-0.0027	
CRIMX	-0.0296		0.0262*
STDPER		0.0056	0.0009*
WHITE		0.0003*	0.0017
INDUT		-0.0041	-0.0059
CMMRC		0.0013*	0.0049
R^2	0.771	0.567	0.315
S	0.1126	0.1866	0.2455

*Not significantly different from zero at .05, 1 tailed test.

[a] This equation was run in double-log form.

Notes to TABLE 2

PRT66 = annual arithmetic mean suspected particulates, 1966, in $\mu g/M^3$/ day.

SUL66 = annual arithmetic mean sulfur dioxide in ppm by volume per seventy-two hours.

TCINC = total current income of family in dollars.

LIVAR = living area of house in square feet.

HSAGE = age of house in years.

MODUR = mortgage duration in years.

LOTSZ = lot size in square feet.

SQIDX = school quality index 1 = highest, 4 = lowest.

INCID = incidental expenses associated with house in dollars.

DISSM = distance to the intersection of State and Madison in tenths of miles.

DLAKE = distance to lake in tenths of miles.

INC66 = median family income in community area, 1966, in dollars.

FRAME = dummy for house type.

DTTYP = dummy for detached.

CRIMX = neighborhood crime rate, an index where the base is the arithmetic mean rate for the City of Chicago.

STDPER = percent standard.

WHITE = percent white.

INPUT = industrial land within one square mile in hundredths of square miles.

CMMRC = commercial land within one square mile in hundredths of square miles

SALEP = sale price.

SITEP = FHA appraiser's estimate of site value.

is used as the dependent variable, were obtained using a specification analogous to Wieand's.

Under each specification, the coefficient of at least one of the pollution measures is significantly negative and the sum of the coefficients is negative. In forms not reported in Table 2 which employ only one pollution measure as an explanatory variable, the coefficient of the included variable is always significantly negative. Evaluated at the means of the relevant variables, the marginal capitalized damages in column 1 are \$474.16 for each additional ten $\mu g/m^3$/ day of suspended particulates plus an additional one ppb per seventy-two hours of sulfur dioxide. In column 2, the analogous damages are \$191.06. A similar computation for column 3 produces marginal capitalized damages of \$.21 per square foot of lot size. This yields an average marginal capitalized damage per residential property of \$869.19, a figure among the highest estimates obtained in all previous studies.

Comparisons of Wieand's results for St. Louis with our own, and comparisons of the results in the first and third columns of Table 2 suggest that Wieand, in general, explains a smaller proportion of the variation in his dependent variable than we explain in the dependent variable we adopt. That is Wieand's specification apparently involves a substantially higher "noise-to-signal ratio" than does that under which we proceed. Other things being equal, one would suppose that it is harder under Wieand's specification than it is under ours to distinguish the relatively minor influence of air pollution on property values from other influences.

FOOTNOTES

1. We assume that the non-negative constraints are not binding.

2. Under common restrictions on $U(\cdot)$, λ is a function of y alone, given p.

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